Economic Policy in Health Care: Sickness Absence and Pharmaceutical Costs

David Granlund

Abstract

This thesis consists of a summary and four papers. The first two concerns health care and sickness absence, and the last two pharmaceutical costs and prices.

Paper [I] presents an economic federation model which resembles the situation in, for example, Sweden. In the model the state governments provide health care, the federal government provides a sickness benefit and both levels tax labor income. The results show that the states can have either an incentive to under- or over-provide health care. The federal government can, by introducing an intergovernmental transfer, induce the state governments to provide the socially optimal amount of health care.

In Paper [II] the effect of aggregated public health care expenditure on absence from work due to sickness or disability was estimated. The analysis was based on data from a panel of the Swedish municipalities for the period 1993-2004. Public health care expenditure was found to have no statistically significant effect on absence and the standard errors were small enough to rule out all but a minimal effect. The result held when separate estimations were conducted for women and men, and for absence due to sickness and disability.

The purpose of Paper [III] was to study the effects of the introduction of fixed pharmaceutical budgets for two health centers in Västerbotten, Sweden. Estimation results using propensity score matching methods show that there are no systematic differences for either price or quantity per prescription between health centers using fixed and open-ended budgets. The analysis was based on individual prescription data from the two health centers and a control group both before and after the introduction of fixed budgets.

In Paper [IV] the introduction of the Swedish substitution reform in October 2002 was used as a natural experiment to examine the effects of increased consumer information on pharmaceutical prices. Using monthly data on individual pharmaceutical prices, the average reduction of prices due to the reform was estimated to four percent for both brand name and generic pharmaceuticals during the first four years after the reform. The results also show that the price adjustment was not instant.

Key Words: vertical fiscal externalities, sickness absence, sickness benefits, health care expenditure, fixed budgets, pharmaceuticals, cost containment, dynamic panel data models, endogeneity, propensity score matching

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David Granlund

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[II] Granlund, D. (2007). The effect of health care expenditure on sickness absence, *Umeå Economic Studies* 701 (revised).

[III] Granlund, D., Rudholm, N. and Wikström, M. (2006). Fixed budgets as a cost containment measure for pharmaceuticals, *The European Journal of Health Economics*, 7, 37-45. (Reprinted with kind permission of Springer Science and Business Media)

[IV] Granlund, D. and Rudholm, N. (2007). Consumer Information and Pharmaceutical Prices: Theory and Evidence, *Umeå Economic Studies* 709.

1 Introduction

This thesis consists of four papers which can be divided into two distinct parts. The first two papers concern the relationship between health care expenditure and absence from work. Paper [I] studies resource allocation problems in a situation where a lower level of government provides health care, the central government provides a sickness benefit and both levels tax labor income. In the second paper, data from a panel of the Swedish municipalities during 1993-2004 was used to study the effect of aggregated health care expenditure on absence.

The last two papers concern cost containment measures taken in the Swedish pharmaceutical markets. The purpose of the third paper was to study whether introducing fixed pharmaceutical budgets for health centers is an effective cost containment measure. The analysis was based on data on individual prescriptions from the county of Västerbotten, in which two health centers were given fixed budgets for pharmaceutical expenditures. Finally, in Paper [IV] a substitution reform in the Swedish pharmaceuticals market was used to study the effects of increased consumer information on pharmaceutical prices.

2 Health care and sickness absence

Health care in Sweden is nearly exclusively publicly provided and private expenditure accounts for only a few percent of total non-dental non-pharmaceutical health care expenditure (Socialstyrelsen, 2006). Today 21 directly elected regional authorities, county councils, are responsible for the public provision. The number of entities has shrunk from 26 a decade ago and a government commission, analyzing the organization of the Swedish public sector, have recently proposed a further reduction of the number of entities responsible for health care provision (SOU 2007:10). The regional authorities finance more than two thirds of their expenditure by proportional labor income taxes. Other important incomesources are grants from the central government, payments from other principals and patients' co-payments (Statistics Sweden; The Swedish Association of Local Authorities and Regions).

In Sweden, the rate of absence from work due to sickness and disability is much higher than in most industrialized countries (OECD, 2005). The rates peaked in 2003 when 5 percent of employee working hours were lost due to sickness absence and 8 percent of the population between the age of 16 and

64 years were on disability pension (Statistics Sweden; The Swedish Social Insurance Agency). One explanation for these high rates is the relatively generous sickness benefit that the central government provides.[1] During the last two decades the compensation levels in the public social insurance system have been above 75 percent of the income from the second day of absence, but with a cap at a certain level of income (Henrekson and Persson, 2004; the Swedish Social Insurance Agency). A second explanation for the high rates is the long waiting times in the health care sector.[2] For example, the waiting time for a primary hip joint operation has exceeded one and a half years in some parts of the country (The Swedish Association of Local Authorities and Regions).

Paper [I] studies the effects of division of power between lower-level governments that provide health care and a central government that provides a sickness benefit. This paper relates to the literature about vertical fiscal externalities. Paper [II] studies the effects of health care expenditure on absence. It primarily relates to the literature evaluating the effects of aggregated health care expenditure and access to health care, but since the effects of health care provided by the Swedish county councils affect the central government's budget constraint, the paper also relates to the empirical literature on vertical fiscal interactions. The existing literature in these fields is reviewed briefly below and thereafter the two papers are summarized.[3]

2.1 Vertical fiscal externalities

The literature about vertical fiscal externalities belongs to the broader group concerning fiscal federalism. Earlier papers in this field focused nearly exclusively on horizontal interactions between lower-level governments "states", and limited the role of the central "federal" government to one that only steps in to resolve the inefficiencies arising from inter-state interactions (Keen, 1998).

Vertical fiscal externalities arise when decisions of governments at one level affect the budget constraint of governments at another level. An early paper

[1] Using Swedish data, Johansson and Brännäs (1998), Johansson and Palme (2002, 2005) and Henrekson and Persson (2004) all found support for that economic incentives affect absence.

[2] Using data from five orthopedic clinics, Statskontoret (2000) found that waiting times affected absence from work.

[3] In addition, Paper [II] relates to the literature on absence. For a review of this literature I refer to Brown and Sessions (1996).

in this field is Cassing and Hillman (1982). They analyzed an Australian case, in which the federal government taxes coal exports and a state government implicitly taxes coal by using its railway monopoly to set excess rail freights on coal transports. Cassing and Hillman demonstrated that total revenues become lower if the federal and state separately try to maximize their revenues, compared to a cooperative solution. Hansson and Stuart (1987), Flowers (1988) and Johnson (1988) were the first to analyze vertical fiscal externalities when federal and state governments impose direct taxes on the same tax base. Flowers provided a result similar to that of Cassing and Hillman, namely that a federation of Leviathan revenue maximizing governments will set taxes above the revenue maximizing level. Johnson demonstrated that a state, by using own taxes to finance redistribution, will reduce income in the state and therefore the residents' federal tax bill. In a situation without migration, Johnson showed that this will make the state's residents prefer redistribution using state taxes as opposed to federal ones.

To avoid inefficiency caused by vertical fiscal externality Hansson and Stuart (1987), Boadway and Keen (1996) and Boadway et al. (1998) proposed the power of taxation to be assigned to only one level of government. Aronsson and Wikström (2001, 2003) demonstrated that this may be unnecessary, since intergovernmental transfer schemes can, in certain situations, induce the correct incentives for the governments.

Only a minority of the literature on vertical fiscal externalities includes expenditure externalities, defined as the effects of a government's expenditure decisions on other governments' budget constraints. Noteworthy exceptions are Dahlby (1996) and Dahlby and Wilson (2003). Dahlby exemplified tax and expenditure externalities and provided general formulas for revenue and expenditure matching grants which can make the governments internalize the externalities. Dahlby and Wilson presented a model where a state government provides a productivity-enhancing public input and both levels of government tax wages and corporate profits. They demonstrated that the vertical externality could be either positive or negative depending on the wage elasticity of the labor demand.

Two empirical papers on fiscal interactions are Besley and Rosen (1998) and Goodspeed (2000); both report evidence of tax interactions between governments. Aronsson et al. (2000) and Andersson et al. (2004) are two studies in this area that use Swedish data. Aronsson et al. found that local expenditure is partly explained by regional expenditure and Andersson et al. report that an increase in the regional tax rate induces municipalities in the region to lower their tax rates. In a recent study using Swiss data Brulhart and Jametti (2006) found that vertical externalities, working for too high taxes, dominate horizontal externalities.

2.2 Health care outcomes

Papers accessing the effects of health care on an aggregated level differ both in terms of dependent and explanatory variables used. Naturally, as dependent variable most studies have used some measure of health status (Nixon and Ulmann, 2006). Infant mortality and life expectancy are often chosen, presumably because that these variables are objective and reasonable easily to measure. The drawback of these measures is that health care has other outcomes than extended life spans that are not captured by these variables, e.g. improvements in health related quality of life. Variables including quality of life are more subjective, less often available, and are therefore not commonly used in the literature. One exception is Miller and Frech (2002). In a cross-section analysis of 18 OECD countries they found significant effects of pharmaceutical expenditure on disability adjusted life expectancy (DALE) but insignificant effects of other health care expenditure.

The measures of health care inputs also differ among studies in this field. Many researchers have used expenditure on health care, often disaggregated to pharmaceutical expenditure and other health care expenditure, as their primary explanatory variables (e.g. Hitiris and Posnet, 1992; Crémieux et al., 1999; Lichtenberg, 2004). Hitiris and Posnet used data from 20 OECD countries over 28 years and found limited effects of health care expenditure on mortality rates. Crémieux et al. found that higher health care expenditure among Canadian provinces reduced infant mortality and increased life expectancy. They explained the limited effect reported by Hitiris and Posnet and others by the inherent heterogeneity associated with cross-country studies. Lichtenberg analyzed time-series of life expectancy in the United States and found that both public health care expenditure and research and development expenditure on pharmaceuticals had positive effects.

Aakvik and Holmås (2006) studied the effect of the number of general practitioners (GPs) on mortality rates in Norwegian municipalities using an instrumental variable estimator. They found no effect of the total number GPs per capita but found a statistically significant negative effect of the number of contracted GPs.

All the studies discussed above are troubled by the context in which decisions regarding health care are taken. For example, the amount of health care provided is partly determined by the perceived need for health care. This results in endogeneity problems, which are not addressed properly in all studies. One study not affected by such problems is Brook et al. (1983). They reported results from a controlled trial in the United States (the Rand Health Insurance Experiment) where families were randomly assigned insurance plans. One group received all their medical care free of charge and, as a consequence, used more than the other groups. The only statistically significant effects on health status were improvements for those with poor vision and for low-income persons with high blood pressure.

2.3 Summary of Papers [I] and [II]

Paper [I]: Sickness absence and health care in an economic federation

In Paper [I] an economic federation model is presented, where the state governments provide health care, the federal government provides a sickness benefit and both levels of government tax labor income. The objective of this paper was twofold. First, to analyze the fiscal externalities facing the state governments and to characterize the differences in health care expenditure and sickness benefit between the centralized and decentralized solutions. Second, to analyze the different ways in which the second best solution can be obtained by changing the governments' responsibilities or policy instruments.

The results show that the states can either have an incentive to under- or over-provide health care depending on, among other things, the wage elasticity of labor supply and the marginal product of expenditure on health care. The federal government can induce the states to increase their health care expenditure by reducing the sickness benefit and the federal tax rate, and vice versa. Whether health care will be under- or over-provided in the decentralized solution depends on the sign of the fiscal externality facing the states, the social costs of financing the sickness benefit and on the slope of the states' reaction functions. By introducing an intergovernmental transfer, the federal government can make the state governments internalize the effects that their decisions will have on the federal government's budget constraint. Moreover, it was proved that the vertical fiscal externality will not vanish by assigning all powers of taxation to the states. This result differs from previous ones presented in the literature and is caused by the fact that the states' decisions in this model directly affect the federal government's expenditure.

The results can be generalized to state financed programs which reduce the number of recipients of federal transfers, for example labor market programs, economic development ventures aimed at reducing poverty and programs aimed at reducing the abuse of federal transfers.

Paper [II]: The effect of health care expenditure on sickness absence

The purpose of Paper [II] was to estimate how aggregated public expenditure on health care affects absence from work due to sickness or disability. To my knowledge, this has not been studied previously.

First, two equations were derived; one that illustrates how absence is affected by health and other variables, and another that shows how health can be affected by public health care expenditure. Then these equations were used in order to specify an empirical model. The analysis was performed using data from a panel of the Swedish municipalities during 1993-2004 and an instrumental variable estimator was used to avoid the potential endogeneity problem caused by absence affecting expenditure on health care.

Public health care expenditure was found to have no statistically significant effect on absence and the standard errors were small enough to rule out all but a minimal effect. This result is robust against changes in model specification and also held when separate estimations were conducted for women and men, and for absence due to sickness and disability.

The main result (that public health care expenditure had a negligible effect on absence) increases the likelihood that general health care is over-provided in Sweden, according to the model in Paper [I]. However, the low effect of health care expenditure may be explained by the Swedish counties' weak incentive to reduce absence which may lead to that a relatively small share of total health care expenditure being focused on reducing absence. Therefore, although aggregated health care expenditure may be over-provided, health care aimed at reducing absence might still be under-provided.

3 Pharmaceutical costs and prices

The Swedish real pharmaceutical costs have doubled in the last 15 years. In 2005 the costs reached 30 billions Swedish crowns and accounted for 13 percent of total health care expenditure (Socialstyrelsen, 2003, 2006). The Swedish development parallels that of many other industrialized countries and has spurred an interest in different methods to contain pharmaceutical costs (Buzzelli et al., 2006).

Since 1993, the price-setting of pharmaceuticals has been unregulated, but for pharmaceuticals to be subsidized, the price charged by the pharmaceutical firms has to be authorized by the Pharmaceutical Benefits Board (PBB). During the past six decades, subsidies have covered a large part of the pharmaceutical costs for Swedish consumers and in 2005 nearly two thirds of the costs were covered by subsidies (Socialstyrelsen, 2006).

In the Swedish pharmaceutical insurance system consumers have to pay all costs below 900 Swedish crowns per year. Then, the subsidy rate gradually increases and all costs exceeding 4,300 Swedish crowns are covered by the insurance (Socialstyrelsen, 2000). There are, however, some exceptions. From 1993 to October 2002 a reference price system was in effect, under which all costs exceeding 110 percent of the price of the least expensive generic substitute also had to be borne by the consumer.[4] On October 1, 2002 a substitution reform came into effect, requiring consumers who refused to switch to the cheapest available substitute to also pay the difference in price between this pharmaceutical and the prescribed one. A key motive behind this reform was to contain pharmaceutical costs (SOU 2000:86). This was also an important motivation for the decentralization of the costs for the pharmaceutical insurance from the central government to the regional governments, that took place in 1998 (Socialstyrelsen, 2000).

In the county of Västerbotten two health centers were, in 2004, given fixed pharmaceutical budgets. Paper [III] studies the effects on pharmaceutical costs of this hardening of previously soft budget constraints. In Paper [IV] the substitution reform is used to study the effects of increased consumer information on pharmaceutical prices. The existing literature in each of these two fields is reviewed briefly below and thereafter the two papers are summarized.

[4] The effects of the reference price system have been analyzed by e.g. Aronsson et al. (2001) and Bergman and Rudholm (2003).

3.1 Soft budget constraints

The term "soft budget constraint" was originally used by Kornai (1979) to denote the budget constraints of enterprises in socialist economies in which deficits were nearly automatically covered by authorities. Primary explanations for the existence of soft budget constraints are asymmetry in information and lack of credible commitments by principals. For example, an investor may when setting up an enterprise, in order to provide management with strong incentives, declare that no additional founds will be provided even if the enterprise would encounter financial difficulties. Thus, the concept is closely related to that of time inconsistency (Kyland and Prescott, 1977). The term is also related to the "ratchet-effect" (Weitzman, 1980), e.g. that a principal may adjust budgets according to history, creating an incentive for agents to overspend to preclude future budget cuts. If the agent realizes that its budget constraint is not hard, this will affect its decisions.

The early literature in the field focused nearly exclusively on state owned enterprises in socialist economies. At present, the concept is also used in analyzes of markets in primarily capitalistic ones. Soft budget constraints clearly apply to the bank sector in most economies, since large banks with severe financial trouble are rarely left without financial assistance and forced to go out of business. The concept can also be used to understand the behavior of local and regional governments, which frequently can rely on being rescued by the central government, and that of poor countries which might obtain international assistance if they become insolvent (Kornai et al., 2003). Further, the budget constraints of nonprofit organizations, such as hospitals and schools, are often soft. For example, Duggan (2000) found clear support for the existence of soft budget constraints for government-owned hospitals in California. Studying the effects of a state-program which made it more profitable to treat the poor, he reported that local government reduced their subsidies to public hospitals by on average one dollar for each additional dollar the hospital earned due to the state-program.

The effects of hardening budget constraint on pharmaceutical costs have been analyzed previously by Whynes et al. (1997). They used cross-sectional data from one English health authority and regressed pharmaceutical costs per patient for each general practitioner (GP). After controlling for five confounding factors, they found that GPs that held their own budgets (and thus had harder budget constraints) had approximately 8 percent lower pharmaceutical costs. However, based on their study it is not possible to separate selection and treatment effects; it was voluntary for GPs to hold their own budgets and it is therefore possible that those who opted for this were those with low initial pharmaceutical costs.

3.2 Consumer Information and Prices

The role of information in economics was largely ignored until Stigler (1961) published his paper on the economics of information. He illustrated that if it is costly to ascertain the most favorable price of a product firms will get market power and set prices above competitive levels. Similarly, Diamond (1971) showed that if information is costly, this could lead to an equilibrium where firms charge monopoly prices. Salop and Stiglitz (1977) presented a model where low-cost stores had higher sales because low search cost individuals actively seek them out, while only high search cost individuals patronized the high cost stores. They also showed that there may exist equilibrium with a single price, which is above the competitive equilibrium price, and that no equilibrium exists under certain circumstances.

Empirical tests of the effects of increased consumer information on prices and market structure are becoming increasingly common. Devine and Marion (1979) collected price information for supermarkets in Ottawa and then published these during a five week period. Compared to the control market, Winnipeg, price levels decreased and consumer satisfaction increased in Ottawa. However, Devine and Marion's paper was later commented on by Lesser and Bryant (1980), who criticized the statistical analysis performed. Recently, empirical papers analyzing the impact of lower search costs through the introduction of internet price comparison sites on price or price dispersion have been published (Baylis and Perloff, 2002; Baye et al., 2004). The findings from these papers show that price dispersion remains even on internet price comparison sites, which according to Bayliss and Perloff could be explained by the division of consumers into two groups, informed and uninformed.

Frank and Salkever (1992) presented a theoretical model for brand name pharmaceutical prices and showed that these will decrease if the share of informed consumers increases. Sorensen (2000) studied how imperfect consumer information affected prices and price dispersion among prescription pharmaceuticals. The data was collected from pharmacies in upstate New York, and the price dispersion among equivalent prescriptions was found to be large. The results also give support to a consumer search cost model, since both markups and price dispersion were significantly lower for frequently purchased pharmaceuticals compared to one-time prescription pharmaceuticals.

3.3 Summary of Papers [III] and [IV]

Paper [III]: Fixed budgets as a cost containment measure for pharmaceuticals

In 2001, two health centers in Västerbotten, Sweden, were given fixed budgets for pharmaceutical expenditure, giving them an incentive to decrease expenditure as they were allowed to keep any surplus (and would be forced to repay any deficit) generated during the year. The purpose of Paper [III] was to analyze if this reform affected the prices and quantities of pharmaceuticals prescribed by physicians working at these health centers.

The analysis was based on data on individual prescriptions from the two health centers and a control group, both before and after the introduction of fixed budgets. Changes in pharmaceutical expenditure were decomposed into three parts; the number of prescriptions, the size of prescriptions (the number of defined daily doses per prescription), and the price of the pharmaceutical. A difference-in-difference extension of propensity score matching was used to study if physicians responded to the budgetary rules by prescribing cheaper medicine or fewer doses of medicine per prescription. This method allowed us to control for observable heterogeneity between patients and for time-invariant unobservable heterogeneity between different health centers.

The results show no systematic effect of the introduction of fixed budgets on either price or quantity per prescription, possibly due to physicians not viewing the fixed budgets to be credible. However, the number of prescriptions in the two health centers with fixed budgets declined relative to the control group after the introduction of the fixed budgets.

Paper [IV]: Consumer Information and Pharmaceutical Prices: Theory and Evidence

In Paper [IV] the impact of increased consumer information on brand name and generic pharmaceutical prices was analyzed both theoretically and empirically. The theoretical results show that an increase in information is likely to reduce the price of brand name pharmaceuticals, while the results regarding generics are less clear.

In the empirical part of the paper, the introduction of the substitution reform in the Swedish pharmaceutical market in October 2002 was used as a natural experiment to examine the effects of increased consumer information on pharmaceutical prices. The main hypothesis to be tested was if the substitution reform, by increasing consumer information about pharmaceutical prices and available generic substitutes, decreased the price of brand name and/or generic pharmaceuticals. In addition, we tested whether the possible price response differed between brand name and generic drugs and studied additional heterogeneity in the reform effect, suggested by the theoretical model. The empirical analysis was based on monthly data on pharmaceuticals sold January 2001 to October 2006 and performed using a Prais-Winsten estimator.

The results from the empirical part of the paper show an average reduction in prices due to the reform of about 4 percent during the period under study, both for brand name- and generic pharmaceuticals. In addition, the results give some support for the reform effect being amplified for pharmaceuticals in markets which had previously been characterized by low levels of consumer information, as well as for pharmaceuticals which prior to the reform had high markups over marginal cost. The results also indicate that the introduction of the reform increased the impact of the number of products on pharmaceutical prices. Finally, the price adjustment was found to be gradual.

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Sickness absence and health care in an economic federation

David Granlund*
Department of Economics
Umeå University
SE-901 87 Umeå
Sweden

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Abstract

This paper addresses vertical fiscal externalities in a model where the state governments provide health care and the federal government provides a sickness benefit. Both levels of government tax labor income and policy decisions affect labor income as well as participation in the labor market. The results show that the vertical externality affecting the state governments' policy decisions can be either positive or negative depending on, among other things, the wage elasticity of labor supply and the marginal product of expenditure on health care. Moreover, it is proved that the vertical fiscal externality will not vanish by assigning all powers of taxation to the states.

Key words: economic federation; moral hazard; vertical fiscal externalities; sickness absence; sickness benefits

JEL classification: H2; H4; H7

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1 Introduction

In Sweden, there is an ongoing debate concerning the high rates of absence from work for health related reasons. The rates have risen significantly over the last decade and, in 2004, 5.2 percent of employee working hours were lost due to sickness absence. At the same time 8.1 percent of the population between the age of 16 and 64 years were on disability pension. One explanation for these high rates is the high benefit levels and the relatively loose regulations regarding sick leave and disability pensions in Sweden. A complementary explanation is the long waiting time for many surgical operations. For example, the waiting time for a primary hip joint operation exceeds one and a half years in some parts of the country.[1]

In Sweden, the levels of the publicly provided sickness benefit and disability pension are decided by central government. However, the county councils are responsible for providing health care. The question that arises, on the basis of this division of power between the two levels of government, is whether it leads to a suboptimal allocation of resources. In this paper an economic federation model is presented, where the state governments provide health care, the federal government provides a sickness benefit and both levels tax labor income.[2] The model resembles the situation in countries such as Finland and Sweden. The purpose of this paper is twofold. First, to analyze the fiscal externalities facing the state governments and to characterize the differences in health care expenditure and level of the sickness benefit between the centralized and the decentralized solutions. Second, to analyze the different ways in which the second best solution can be obtained by changing the governments' responsibilities or policy instruments. In order to focus on vertical externalities, this paper ignores the horizontal ones.

The paper primary relates to the literature concerning vertical fiscal externalities but also to the literature relating to absence from work. Brown and Sessions (1996) provide a survey of the literature in the latter field. Most of

[1] The data are obtained from Statistics Sweden (www.scb.se), The Swedish Social Insurance Agency (www.forsakringskassan.se) and The Swedish Association of Local Authorities and Regions (sas.lf.se).

[2] In this paper "federal government" is used to denote the central government and "state government" is used to denote the lower level of government, like state, regional or local level government.

the literature in this area is empirical and includes estimations of the effects of working conditions, stipulated work time, job satisfaction and overtime, on work absence. Also, the literature provides strong support for the idea that economic incentives impact upon absence behavior.

Vertical fiscal externalities when federal and state governments impose taxes on the same tax base were first considered by Hansson and Stuart (1987), Flowers (1988) and Johnson (1988). Flowers showed that a federation of Leviathan revenue maximizing governments will end up on the downward sloping section of the Laffer curve. Johnson showed that residents of a state prefer to redistribute using state taxes as opposed to federal taxes since this, by reducing the money income in the state, reduces the state's federal tax bill, meaning that some of the cost of the redistribution will be born by residents of other states. To deal with the vertical fiscal externality Hansson and Stuart (1987), Boadway and Keen (1996) and Boadway et al. (1998) proposed that the power of taxation be assigned to only one level of government. Aronsson and Wikström (2001, 2003) showed that intergovernmental transfer schemes can, in certain situations, induce the correct incentives, making it unnecessary to restrict the taxation power to one level of government.[3]

The majority of the literature in this field focuses on the externality that arises from the co-occupancy of a common tax base, as opposed to expenditure externalities, and has come to the conclusion that the vertical externality is negative.[4] An interesting exception is Dahlby and Wilson (2003). In their model a state government provides a productivity-enhancing public input and both levels of government tax wages and corporate profits. They show that the vertical externality can be either positive or negative depending on the wage elasticity of the labor demand. In this paper, a different kind of expenditure externality is examined. Instead of providing a productivity-enhancing public input, the states in this paper provide a private good, health care, which affects the fraction of the population that is able to work. The state's decisions not only affect the federal government's tax revenues but also affect its sickness benefit expenditure. The results show that the states will have an incentive to either under-provide or over-provide health care depending on, among other things,

[3] Keen (1998) reviewed the literature on vertical fiscal externalities.

[4] Tax and expenditure externalities are here defined as the effects of a government tax and the effects of expenditure decisions, respectively, on other governments' budget constraints.

the wage elasticity of the labor supply and the marginal product of expenditure on health care at the equilibrium. The federal government can induce the states to increase (reduce) their expenditure on health care, by reducing (increasing) the sickness benefit and the federal tax rate. Whether health care will be under- or over-provided in the decentralized solution depends on the sign of the fiscal externality facing the states, the social costs of financing the sickness benefit and on the slope of the states' reaction functions. The federal government is able to achieve the second best solution if it is given the possibility of deciding an intergovernmental transfer.

In comparison to earlier studies, this paper contributes to the literature by letting the state governments provide a good that affects the share of the population that works and hence the federal government's transfers to those not working. The results can be generalized to state financed programs which reduce the number of recipients of federal transfers, for example labor market programs, economic development ventures aimed at reducing poverty and programs aimed at reducing the abuse of federal transfers. Another contribution is that this paper shows that vertical externalities will not vanish by assigning all powers of taxation to the state governments if federal expenditure depends directly on the decisions taken by the states.

The rest of the paper is organized as follows: In section 2 the model is set up and the decision rules for a unitary central government, which will serve as a benchmark in the following, are derived. Section 3 presents the decentralized solution. First, in subsection 3.1 the policy decisions that the state governments will take if they act as Nash competitors to one another and towards the federal government are examined and compared with the policy decisions of the unitary central government. Here the states' reactions to changes in the federal governments' policy variables are also derived. The following subsection describes the policy decisions taken by the federal government, when it acts as a first mover. Different ways of implementing the second best solution are discussed in subsection 3.3. Finally, in section 4 the paper's conclusions are presented.

2 The model

The federation consists of N identical states, which are small relative to the federation. Each state is populated by a continuum of residents normalized to unity. The federal government pays a sickness benefit to individuals on sick leave and the state governments are responsible for providing health care. Following Boadway and Keen (1996), both levels of government are assumed to finance their expenditure via a proportional labor income tax and to balance their budgets. Residents are assumed to be immobile between states and benefit spillovers from health care are assumed not to exist. These assumptions allow us to focus on the vertical fiscal externalities and ignore possible horizontal externalities. Since the states are identical, no index for states is used and the analysis is focused on a representative state, and due to the lack of horizontal externalities N is normalized to one.

The utility function of an individual is written, $U_i = u(c, l, h) - g(m_i)$, where $i \in [0, 1]$, c is private consumption, l labor supply and h health state. Preferences are identical and the utility is increasing in c, decreasing in l and strictly concave in c and l. Individuals have no access to capital markets, so they consume all of their income. There are two health states; healthy, h^h, and sick, h^s, where the first yields higher utility for given levels of c and l. $m_i \in [\underline{m}, \overline{m}]$ is a moral parameter that varies continuously among individuals and has a rectangular distribution. $-g(m)$ shows the non-pecuniary disutility from pretending to be sick, when healthy. A high value of m depicts a high moral and g is assumed to be increasing and concave in m.

By letting individuals differ in respect to a moral parameter, I allow for the possibility of a positive fraction of mimickers in the respective solutions. The motive for doing so is primarily to add realism to the model. In reality individuals are heterogeneous and it is, therefore, sub-optimal for a government with limited instruments to form policies so that no one mimics another. Second, letting the individuals be heterogeneous in this respect makes it possible to demonstrate how the presence of mimickers affects the vertical externalities. An alternative to including moral in this form is to allow for heterogeneity with respect to the value of consumption or leisure, or to allow for a continuum of health states. Additive separability in g is assumed for simplicity.

The health function, equation (1) below, which is the steady state solution to the dynamic equation,[5] describes the relationship between the sickness rate

[5] The dynamic equation is written

$$\dot{a} = d(1 - a) - ra - f(e).$$

(a) and the public health care expenditure per capita (e).

$$a(e) = \frac{d - f(e)}{r + d}. \tag{1}$$

Healthy individuals turn sick at the exogenous rate d and sick individuals recover by themselves at the exogenous rate r. Both d and r take values between zero and one. $f(e)$ denotes the rate at which the population recovers as a result of medical treatment.[6] Health care is assumed to be exclusively financed by the governments and patients are assumed to experience no disutility by treatment. $f' > 0$ and $f'' < 0$ are assumed, stating that health care expenditure have a positive effect on the number of recoveries but that the effect of expenditure is decreasing.[7]

The justification for the assumption that health care is exclusively financed by the governments is two-fold. First, the purpose of this paper is to study the interactions between different levels of government when health care and sickness benefits are publicly provided. Second, incorporating motives for the public provision, instead of just assuming it, would complicate the model without contributing to the understanding of the problem. Among the possible motives for the public provision of health care and sickness benefit, distributional objectives and adverse selection can be mentioned.

In the model presented here, all sick individuals will be on sick leave. Healthy individuals can either work (workers) or be on sick leave (mimickers). Let $s(e) = a(e) + \hat{a}(1 - a(e))$, where $s(e)$ is the absence rate[8] and $\hat{a}$ is the fraction of healthy individuals who are mimickers. Both levels of government are assumed to know the size of $f(e)$ for all possible levels of e as well as the value of the parameters r and d; $a(e)$, $s(e)$ and $\hat{a}$ are also assumed to be observable by them. However, the governments can not observe an individual's morals or whether the individual is sick or a mimicker. In reality, governments that provide sickness

[6]Since the population is constant, this is a way of modeling the number of treated individuals as a monotone function of health care expenditure, but does not imply that that medical treatment is applied to the whole population. An alternative would be to let $f(e)$ denote the rate at which the sick fraction of the population recovers as a result of medical treatment. However, this would imply that the cost of treating a given share of the sick is independent of the number of sick individuals, which is clearly unrealistic.

[7]Expenditure on health care might affect the sickness rate by, for example, leading to a reduction in waiting times or improved procedures.

[8]In this paper, the absence rate can be interpreted to include not only the sickness absence rate among employees, but also the rate of individuals on disability pension.

benefits often attempt to distinguish mimickers from the sick by requiring a doctor's certificate as a prerequisite for receiving a sickness benefit, at least for long-term sick leave. However, physicians cannot distinguish perfectly between a sick individual and a mimicker, and even if they can, the incentives to reveal a mimicker is often limited (see e.g. Shortell (1998) for a discussion of physician's multiple accountabilities). The governments are, therefore, unlikely to be able to identify all mimickers. One could model this by introducing some probability for a mimicker to be detected and perhaps a penalty if detected. This would reduce the problem with mimickers, but would not change the general results. Further as introducing this factor would complicate the model, the extreme case where governments have no possibility of detecting a mimicker is chosen.

The private agents are assumed to make their choices concerning labor supply and sick leave after the governments' policies have been proclaimed. The workers choose consumption and labor supply to maximize their utility, subject to their budget constraints

$$c = w(1-t-T)l,$$

where t and T are tax rates imposed by the state and federal governments, respectively, and w is the exogenously given real wage rate. The outcome of a worker's optimization problems will be $c = c(w(1-t-T))$ and $l = l(w(1-t-T))$. For individuals on sick leave $c = B$, where B is the sickness benefit. The indirect utility function for workers, and the conditional indirect utility functions for mimickers and sick individuals, respectively, are written

$$V^h = v(w(1-t-T), h^h),$$

$$\widehat{V_i^h} = \widetilde{v}(B, h^h) - g(m_i),$$

$$V^s = \widetilde{v}(B, h^s),$$

where tilde indicates that the indirect utilities for mimickers and sick individuals are conditioned on the labor supply being fixed at zero. Healthy individuals will be on sick leave if

$$\widetilde{v}(B, h^h) - g(m_i) \geq v(w(1-t-T), h^h), \tag{2}$$

therefore $\hat{a} = \hat{a}(t, T, B)$. Equation (2) tells us that for the last individual who chooses to be a mimicker $m_i = g^{-1}(x)$, where $x = \widetilde{v}(B, h^h) - v(w(1-t-$

$T), h^h)$. The rectangular distribution of m, together with the concavity of $g(m)$, is sufficient to prove the following results:

$$\begin{aligned}
\frac{\partial \hat{a}}{\partial t} &= \frac{\partial \hat{a}}{\partial T} = g^{-1\prime}(x) v'(w(1-t-T), h^h) w > 0 \\
\frac{\partial \hat{a}}{\partial B} &= g^{-1\prime}(x) \widetilde{v}'(B, h^h) > 0 \\
\frac{\partial^2 \hat{a}}{\partial t \partial T} &= \frac{\partial^2 \hat{a}}{\partial t^2} = \frac{\partial^2 \hat{a}}{\partial T^2} = -g^{-1\prime}(x) v''(w(1-t-T), h^h) w^2 \\
&\quad + g^{-1\prime\prime}(x) [v'(w(1-t-T), h^h) w]^2 > 0 \\
\frac{\partial^2 \hat{a}}{\partial t \partial B} &= v'(w(1-t-T), h^h) w g^{-1\prime\prime}(x) \widetilde{v}'(B, h^h) \geq 0 \\
\frac{\partial^2 \hat{a}}{\partial B^2} &= g^{-1\prime}(x) \widetilde{v}''(B, h^h) + g^{-1\prime\prime}(x) [\widetilde{v}'(B, h^h)]^2 \gtreqless 0.
\end{aligned}$$

Ranking individuals by increasing morals and assuming the welfare objective to be utilitarian[9], permits the governments' objective function to be written

$$\begin{aligned}
&(1-a(e))(1-\hat{a})\ v(w(1-t-T), h^h) + a(e)\widetilde{v}(B, h^s) \\
&+(1-a(e)) \int_{i=0}^{\hat{a}} [\widetilde{v}(B, h^h) - g(m_i)] di. \qquad (3)
\end{aligned}$$

2.1 Centralized policy decisions

In this section the policy decisions that a central government in a unitary nation would take, in the absence of any fiscal responsibility of any lower level of government, are derived. These policy decisions will later be used as a benchmark against which the decentralized policy decisions will be compared. When all policy decisions are made by the central government there is no need to distinguish between the federal and state tax rates. It is therefore assumed that the

[9] One could allow the governments to differently weight the utility of workers, sick individuals and mimickers. Perhaps the most reasonable alteration would be to assign a lower weight to mimickers, since they abuse the system (see e.g. Sandmo (1981) for a discussion about this). This would affect the decisions taken both in the centralized and decentralized setting, but would not contribute to the understanding of the vertical fiscal externalities analyzed in this paper. Since this also would expand the notation and require more extensive explanations of the equations to follow, this is not done. However, the main effects of assigning a lower weight to mimickers will be mentioned in the paper.

government chooses a single rate $\tau = T + t$. The decision variables are τ, B and e. The second order conditions for maximums are assumed to be fulfilled and the solution is assumed to imply positive levels for all variables. The optimization problem of the central government in a unitary nation coincides with the social optimization problem and can be written

$$\underset{\tau,B,e}{\text{Max}}\ (1-a(e))(1-\hat{a})\ v(w(1-\tau),h^h)+a(e)\widetilde{v}(B,h^s)$$

$$+(1-a(e))\int\limits_{i=0}^{\hat{a}}[\widetilde{v}(B,h^h)-g(m_i)]di$$

subject to the budget constraint

$$(1-a(e))(1-\hat{a})\tau wl-(a(e)+\hat{a}(1-a(e)))B-e=0,$$

where $a(e)$, $\hat{a}$ and l are defined as before. The Lagrangian becomes

$$\begin{aligned} L &= (1-a(e))(1-\hat{a})\ v(w(1-\tau),h^h)+a(e)\widetilde{v}(B,h^s) \\ &+ (1-a(e))\int\limits_{i=0}^{\hat{a}}[\widetilde{v}(B,h^h)-g(m_i)]di \\ &+ \gamma\{(1-a(e))(1-\hat{a})\tau wl-(a(e)+\hat{a}(1-a(e)))B-e\}, \end{aligned}$$

where γ is the Lagrangian multiplier, which at the optimum can be interpreted as the marginal cost of public funds. Given that the condition in equation (2) results in equally high utilities for a worker and for the last individual who chooses to be a mimicker, the first order conditions can be written

$$\begin{aligned} \tau &: -(1-\hat{a})v'(w(1-\tau),h^h)w \\ &\quad +\gamma\{(1-\hat{a})(wl-\tau w^2 l')-(\tau wl+B)\frac{\partial\hat{a}}{\partial\tau}\}=0 \end{aligned} \tag{4}$$

$$\begin{aligned} B &: a(e)\widetilde{v}'(B,h^s)+(1-a(e))\hat{a}\widetilde{v}'(B,h^h) \\ &\quad -\gamma\{a(e)+\hat{a}(1-a(e))+(1-a(e))(\tau wl+B)\frac{\partial\hat{a}}{\partial B}\}=0 \end{aligned} \tag{5}$$

$$e \quad : \quad -\left[(1-\hat{a})V^h - V^s + \int_{i=0}^{\hat{a}} \widehat{V_i^h} di\right] a'(e)$$
$$-\gamma\{(1-\hat{a})(\tau wl + B)a'(e) + 1\} = 0. \qquad (6)$$

Note that the expression in curly brackets in equation (4), except $B\frac{\partial \hat{a}}{\partial \tau}$, multiplied with $(1 - a(e))$ represents the slope of the so called 'Laffer curve'.[10] Since $v'(w(1-\tau), h^h) > 0$, $(1 - a(e)) > 0$ and $B\frac{\partial \hat{a}}{\partial \tau} > 0$, equation (4) implies that the tax revenue is a non-decreasing function of the tax rate in the unitary solution. Throughout this paper, l' is assumed to be strictly positive, stating that the substitution effect always dominates the income effect.

Equation (5) shows that the marginal benefit of consumption, for the average person on sick leave, will be higher than the marginal cost of public funds. The reason for this is that the government holds back the level of the sickness benefit, since it will affect the fraction that works.

The first row of equation (6) shows the marginal benefit of expenditure on health care. Since a fraction of those who will be treated will become mimickers, the expression consists of a weighted sum of workers' utility and mimickers' utility, minus the utility of sick individuals. Throughout this paper, the utility of the average healthy individual is assumed always to exceed the utility of a sick individual. $-(1-\hat{a})(\tau wl + B)a'(e)$ is a 'health budget feedback effect' that shows the extra tax revenues and public savings that expenditure on health care will cause. Let τ^*, B^*, e^*, denote the second best tax and expenditure policies.[11]

[10] In addition to the standard terms in the Laffer curve, which describes the relationship between tax revenues and the tax rate through the workers labor supply decision, the expression includes the term $\tau wl\frac{\partial \hat{a}}{\partial \tau}$, which multiplied with $(1-a(e))$ describes how the tax revenues are affected by the tax rate through its effect on the number of workers.

[11] If mimickers were assigned the weight ζ, $0 \leq \zeta < 1$, in the social welfare function, this would result in lower levels of τ^*, B^*, e^* compared to the case above, when the weight is unity. This can be seen by inserting ζ before the mimickers' utilities and marginal utilities in the first order conditions and adding the terms $[\zeta\widehat{V_{\hat{a}}^h} - V^h]\frac{\partial \hat{a}}{\partial \tau}$ and $[\zeta\widehat{V_{\hat{a}}^h} - V^h]\frac{\partial \hat{a}}{\partial B}$ to equations (4) and (5), respectively. Here $\widehat{V_{\hat{a}}^h}$ denotes the utility of the marginal mimicker. Since equation (2) gives that $\widehat{V_{\hat{a}}^h} = V^h$, these new terms will be negative and therefore work for a lower tax rate and a lower sickness benefit. The latter would also be reduced compared to the case when $\zeta = 1$, since part of the utility of the benefit goes to mimickers. Further, the government's incentive to cure individuals would be reduced, since some of the cured ones would become mimickers.

3 Decentralized policy decisions

This section begins with a description of the optimization problems facing the state and federal governments, and then different methods of implementing the second best resource allocation are discussed. The decisions of the federal government are important for all states and therefore the federal government is assumed to act as a first mover, committing to its policies before the states and anticipating their effects on the states' decisions. The consequences for the federal government of the actions taken by a single small state are minor and the states are therefore assumed to act as Nash competitors towards it. As mentioned, no interactions between the states are assumed to exist and hence the states take the decisions by the other states as given.

3.1 The state governments

Since all states are identical, we can focus on a representative state. The state government chooses t and e to maximize its objective function subject to its budget constraint. The optimization problem can be written

$$\underset{t,e}{\text{Max}}\ (1-a(e))(1-\hat{a})\ v(w(1-t-T),h^h)+a(e)\widetilde{v}(B,h^s)$$

$$+(1-a(e))\int\limits_{i=0}^{\hat{a}}[\widetilde{v}(B,h^h)-g(m_i)]di$$

s.t.

$$(1-a(e))(1-\hat{a})twl-e=0,$$

where $a(e)$, $\hat{a}$ and l are defined above. The first order conditions can be written

$$\begin{aligned} t &: \quad -(1-\hat{a})v'(w(1-t-T),h^h)w \\ &\quad +\gamma^s\{(1-\hat{a})(wl-tw^2l')-twl\frac{\partial\hat{a}}{\partial t}\}=0 \\ e &: \quad -\left[(1-\hat{a})V^h-V^s+\int\limits_{i=0}^{\hat{a}}\widehat{V_i^h}di\right]a'(e) \\ &\quad -\gamma^s\{(1-\hat{a})twla'(e)+1\}=0, \end{aligned}$$

The decentralized solution is affected correspondingly.

where γ^s denotes the Lagrangian multiplier, which at the optimum can be interpreted as the state's perceived marginal cost of public funds.

Comparing these conditions with those for the unitary nation, we see that they neglect the effect that the policy decisions will have on the federal budget. The tax externality works for a too high t and e if $T > 0$ or $B > 0$, but the opposite is true for expenditure externality. Which effect that dominates is inconclusive, given the assumptions made. Even if the federal tax rate is zero, the states may have suboptimal incentives since they neglect the effect that their decisions will have on the federal government's expenditure for sickness benefit.

To determine under which conditions the state will have an incentive to under- and over-provide health care, an expression for the vertical fiscal externality has to be derived. Given that the federal government's budget constraint, R_f, can be written

$$R_f = (1 - a(e))(1 - \hat{a})Twl - (a(e) + \hat{a}(1 - a(e)))B, \tag{7}$$

the effect of a state's expenditure on health care on R_f can be written

$$\frac{dR_f}{de} = -(1 - a(e)) \left[(1 - \hat{a})Tw^2 l' + (Twl + B)\frac{\partial \hat{a}}{\partial t} \right] \frac{dt}{de} - (1 - \hat{a})(Twl + B)a'(e) \tag{8}$$

$$\frac{dt}{de} = \frac{1 + (1 - \hat{a})twla'(e)}{(1 - a(e)) \left[(1 - \hat{a})(wl - tw^2 l') - twl\frac{\partial \hat{a}}{\partial t} \right]}, \tag{9}$$

where $\frac{dt}{de}$ originates from the state's budget constraint. The first part of the equation describes the indirect effect that the state's health care expenditure have on the federal budget constraint, through its relationship with the state's tax rate given by the state's budget constraint. This indirect effect works through the effect of the tax rate on the workers' labor supply and the rate of mimickers. The last term shows how the state's health care expenditure influence the federal budget constraint by changing the share of the population that works and therefore not only altering the federal tax income but also the federal government expenditure on sickness benefit. A positive vertical fiscal externality, $\frac{dR_f}{de} > 0$, means that the state will have an incentive to under-provide health care.

Given that the federal budget constraint can be written

$$(1 - s(e))Twl = s(e)B. \tag{10}$$

and multiplying by $\frac{de}{dt}$ gives us Proposition 1 below, where $\eta = (\partial l/\partial w)(w/l)$ is the wage elasticity of the labor supply among workers and the term $\frac{de}{dt}$ is the inverse of equation (9). Given the states' first order conditions, $\frac{de}{dt}$ is positive.

Proposition 1 *In a decentralized setting where the state governments act as Nash competitors and given any level of the federal government's tax rate and the sickness benefit, the states will under-provide (over-provide) health care if*

$$s(e)\eta + \frac{1}{1-\hat{a}}\frac{\partial \hat{a}}{\partial t} < (>) \frac{-1}{1-a(e)}a'(e)\frac{de}{dt}. \tag{11}$$

The proposition shows that health care will be under-provided by the states for any given level of B and T if the tax externality (the left hand side of equation (11)) is dominated by the expenditure externality (the right hand side). That is, health care will be under-provided if an increase in the states' tax rates will have a smaller impact on the federal government's budget constraint than the expenditure on health care that the tax increase can finance.

Proposition 1 shows that the less sensitive healthy individuals' total labor supply is to taxes, the more likely health care is to be under-provided by the state government. The workers' labor supply elasticity will be more important the higher the fraction of individuals on sick leave. This may seem counter-intuitive at first, but is explained by the form of the federal governments' budget constraint. Equation (10) shows that a high value of $s(e)$ implies that federal tax revenues from a worker have to be large in comparison to the level of the sickness benefit. Therefore, changes in the labor supply will cause relatively large effects on the federal tax revenues. Other things being equal, health care is less likely to be under-provided the higher the share of mimickers. The intuition is that if a large share of the healthy individuals is represented by mimickers, the fiscal reason for treating the sick becomes less important, reducing the expenditure externality.[12]

Equation (11) can be rewritten using the state government's decision rule. Maximizing the state's objective function with respect to e, and letting t be

[12] Proposition 1 holds even if the social weights assigned to the utilities of mimickers and other groups are changed, since these not directly affect the governments' budget constraints.

defined subsequently by the state's budget constraint, gives the first order condition

$$-\left[(1-\hat{a})V^h - V^s + \int\limits_{i=0}^{\hat{a}} \widehat{V_i^h} di\right] a'(e) - (1-a(e))(1-\hat{a})v'(w(1-t-T), h^h)w\frac{dt}{de} = 0. \tag{12}$$

This allows equation (11) to be written as

$$s(e)\eta + \frac{1}{(1-\hat{a})}\frac{\partial \hat{a}}{\partial t} < (>)(1-\hat{a})\frac{v'(w(1-t-T), h^h)w}{(1-\hat{a})V^h - V^s + \int\limits_{i=0}^{\hat{a}} \widehat{V_i^h} di}. \tag{13}$$

The denominator of the right hand side of equation (13) shows the indirect utilities for individuals in different states. If the redistribution in the society is extensive, the utility of sick individuals will approach the utility of the average healthy individual. In this case, health care will be under-provided, since the expression on the right hand side of equation (13) will approach infinity. The intuition is straight forward. If the average individual only experiences a very little increase in utility by getting treated, there is little incentive for the state to provide health care.

Before we continue to the federal level, it is helpful to analyze how the state government will react to the federal government's decisions. Below, I describe how the state will alter its expenditure on health care in response to changes of T and B, respectively, and letting the state's tax rate be defined subsequently by its budget constraint. In order to do so, equation (12) is differentiated with respect to e and T or B, holding the other federal governments' policy instruments fixed. This gives the following expressions

$$\frac{de}{dT} = -\frac{\begin{array}{c} a'(e)(1-\hat{a})v'w + (1-a(e))(1-\hat{a})v''w^2\frac{dt}{de} \\ +(1-a(e))v'w\frac{dt}{de}\frac{\partial \hat{a}}{\partial T} - (1-a(e))(1-\hat{a})v'w\frac{\partial(dt/de)}{\partial T}\end{array}}{\delta} \tag{14}$$

$$\frac{de}{dB} = -\frac{\begin{array}{c} a'(e)\left[\frac{\partial \tilde{v}(B,h^s)}{\partial B} - \hat{a}\frac{\partial \tilde{v}(B,h^h)}{\partial B}\right] \\ +(1-a(e))v'w\frac{dt}{de}\frac{\partial \hat{a}}{\partial B} - (1-a(e))(1-\hat{a})v'w\frac{\partial(dt/de)}{\partial B}\end{array}}{\delta}, \tag{15}$$

where δ is the second derivative of the state's optimization problem with respect to e.

As will be demonstrated, both the numerators and the denominators of equations (14) and (15) are negative, meaning that the state will react to an increased federal tax rate or sickness benefit by reducing its expenditure on health care.

The denominators, δ, are negative given that the state's objective function is concave in e.

The first term in the numerator of equation (14) shows that an increased federal tax rate will reduce the workers' utility, reducing the state's incentive to cure sick individuals. The other terms demonstrate how the state's perception of the marginal costs of public funds is affected by the federal tax rate. An increased federal tax rate increases the workers' marginal benefit of income and reduces the amount of health care that can be financed by a given tax rate, since it reduces the total number of hours worked. Both these effects tend to increase the state's perception of the marginal costs of public funds when the federal tax rate is increased. However, a higher federal tax rate also reduces the number of workers that are affected by the tax rate, by increasing the share of mimickers. This effect goes in the opposite direction to the other two. In the appendix it is demonstrated that this effect is dominated by the effect that works through changing the relation between t and e.

The first term in the numerator of equation (15) shows that an increased sickness benefit will increase the utility of people on sick leave. This will reduce the state's incentive to cure sick individuals, if $\frac{\partial \widetilde{v}(B,h^s)}{\partial B} > \hat{a}\frac{\partial \widetilde{v}(B,h^h)}{\partial B}$. This will be the case unless consumption and health are sufficiently strong complements, and the number of mimickers is sufficiently large. This is extremely unlikely and the first term will therefore be assumed to be negative.[13] The other two terms illustrate how the state's perception of the marginal costs of public funds is affected by the sickness benefit. As demonstrated in the appendix these terms will be jointly negative.

To conclude, given the assumptions made, an increase in the federal tax rate or the sickness benefit reduces the state's incentive to cure sick individuals and increases its perception of the marginal costs of public funds, causing it to

[13] Empirical estimates reported in the literature indicate that consumption and health are complements, but not strong enough to make the first term in equation (15) positive. Viscusi and Evans (1990) estimate the marginal utility of consumption when ill to be 77.3 per cent of that when well. The corresponding estimates reported in Gilleskie (1998) for acute illness are 58.2 and 15.6, depending on the type of illness. Since the rate of mimickers is indeed below these figures, these estimates support the assumption that the first term in equation (15) is negative.

reduce its expenditure on health care. Due to the balanced budget constraint, this might imply a reduction of the state's tax rate. However, a reduction of the expenditure on health care does not by itself guarantee that the state's tax rate will also be reduced, since a change in anyone of the federal government's policy variables alters the relationship between the state's policy variables. How the state will adjust its tax rate, when the federal government changes its policy choices, can be derived and interpreted in the same manner as for the state's adjustment of its expenditure on health care. In the appendix it is demonstrated that $\frac{dt}{dT} < 0$ and $\frac{dt}{dB} < 0$.

3.2 The federal government

The federal government acts as a first mover and chooses T and B to maximize its objective function subject to its budget constraint, the private agents' responses and the states' reaction functions just described. The problem can be written

$$\underset{T,B}{\text{Max}}\ (1-a(e))(1-\hat{a})\ v(w(1-t-T),h^h)+a(e)\widetilde{v}(B,h^s)$$

$$+(1-a(e))\int\limits_{i=0}^{\hat{a}}[\widetilde{v}(B,h^h)-g(m_i)]di$$

s.t.

$$(1-a(e))(1-\hat{a})Twl-(a(e)+\hat{a}(1-a(e)))B=0,$$

where $e=e(B,T)$, $t=t(e(B,T),B,T)$ and where $a(e)$, $\hat{a}$ and l are defined as before. Letting γ^f denote the Lagrangian multiplier, which at the optimum can be interpreted as the federal government's perceived marginal cost of public funds, the federal government's first order conditions can be written

$$\begin{aligned} T\ &:\ -(1-a(e))(1-\hat{a})v'(w(1-t-T),h^h)w \\ &\quad +\gamma^f(1-a(e))\left[(1-\hat{a})(wl-Tw^2l')-(Twl+\ B)\frac{\partial\hat{a}}{\partial T}\right]+\delta_T=0 \\ B\ &:\ a(e)\widetilde{v}'(B,h^s)+(1-a(e))\hat{a}\widetilde{v}'(B,h^h) \\ &\quad -\gamma^f\left\{a(e)+\hat{a}(1-a(e))+(1-a(e))(Twl+\ B)\frac{\partial\hat{a}}{\partial B}\right\}+\delta_B=0, \end{aligned}$$

where

$$\begin{aligned}
\delta_T &= [-(1-a(e))(1-\hat{a})v'w \\
&+\gamma^f(1-a(e))\left[(1-\hat{a})(-Tw^2l') - (Twl + B)\frac{\partial\hat{a}}{\partial t}\right]]\frac{dt}{dT}|_{de=0} \\
&+\gamma^f Z\frac{de}{dT},
\end{aligned}$$

$$\begin{aligned}
\delta_B &= [-(1-a(e))(1-\hat{a})v'w \\
&+\gamma^f(1-a(e))\left[(1-\hat{a})(-Tw^2l') - (Twl + B)\frac{\partial\hat{a}}{\partial t}\right]]\frac{dt}{dB}|_{de=0} \\
&+\gamma^f Z\frac{de}{dB},
\end{aligned}$$

$$\frac{dt}{dT}|_{de=0} = -\frac{(1-\hat{a})(-tw^2l') - twl\frac{\partial\hat{a}}{\partial T}}{(1-\hat{a})(wl - tw^2l') - twl\frac{\partial\hat{a}}{\partial t}} > 0,$$

$$\frac{dt}{dB}|_{de=0} = \frac{twl\frac{\partial\hat{a}}{\partial B}}{(1-\hat{a})(wl - tw^2l') - twl\frac{\partial\hat{a}}{\partial t}} > 0,$$

$$Z = -(1-a(e))\left[(1-\hat{a})Tw^2l' + (Twl + B)\frac{\partial\hat{a}}{\partial t}\right]\frac{dt}{de}$$

$$-(1-\hat{a})(Twl + B)a'(e).$$

Here, δ_T and δ_B represent the indirect effects of the federal government's decision variables on the Lagrangian, via the reaction function for the states' expenditure on health care and their budget constraints. The first two rows in the expressions for δ_T and δ_B, respectively, describe the effect that the federal government's decision variables have on the Lagrangian through their effects on the state's tax bases. These terms work in favor of lower T and B, respectively, since this will reduce the states' tax rates for any given level of e. Z relates to the vertical externality facing the state governments, described in equation (8). If the states have an incentive to under-provide health care, Z will be positive, working in favor of lower T and B, and vice versa.

A special case appears when the tax and expenditure externalities described in equation (8) compensate each other exactly. Z will then equal zero and the first order conditions will be identical to the first order conditions for the unitary nation.[14] In this case the federal government has enough instruments at its disposal to obtain the unitary nation optimum. By setting $B = B^*$ the federal government will induce the state governments to set $e = e^*$ and the budget constraints gives $(t + T) = \tau^*$.

In general, the externalities facing the state governments will not cancel out and the federal government will not be able to simultaneously achieve both B^* and e^*, given that B and T are its only two policy variables. Instead, the federal government is left to choose a point on the state governments' reaction functions. The different situations are illustrated in Figure 1, where the problem is reduced to the state governments choosing e and the federal government choosing B, letting t and T be defined subsequently by the respective budget constraints. $Z < 0$ and $Z > 0$ illustrate what the states' reaction functions can look like if the states have an incentive to over-provide and under-provide health care, respectively. $Z = 0$ is an illustration of the case where the tax- and expenditure externalities cancel out exactly. U_1 and U_2 ($U_1 < U_2$) illustrate what the federal government's indifference curves in the $B - e$ plane might look like given the relationship between these variables and the tax rates.

Figure 1 illustrates a setting where the federal government will choose a point in the North East quadrant if the states have incentives to over-provide health care. By setting B above B^* the federal government has induced the states to reduce their expenditure on health care compared to point 1. However, e is still above e^* and the total tax rate exceeds that of the unitary solution.

Given the general form of the model, we can not conclude that the solution will be in this quadrant. If the high state tax rates, associated with high expenditure on health care, result in sufficiently high social cost for financing the sickness benefit and if the slope of the states' reaction functions is sufficiently flat, the federal government will choose a point in the North West quadrant. In the opposite case the federal government will choose a point in the South East quadrant, which means that, by selecting a high enough sickness benefit, it will induce the states to choose e below e^*, despite their incentives to choose a too high level of e for any given level of B.

[14] To see this, use that $\frac{\partial \hat{a}}{\partial t} = \frac{\partial \hat{a}}{\partial T}$, insert the expressions for $\frac{dt}{dT}|_{de=0}$ and $\frac{dt}{dB}|_{de=0}$, respectively, in the first order conditions and rearrange.

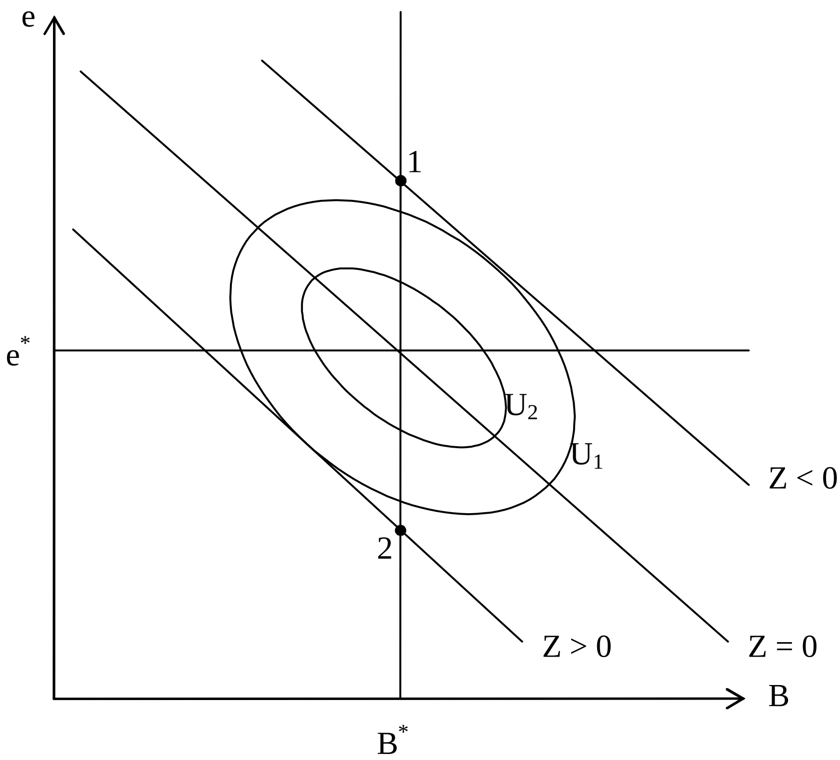

Figure 1. Illustration of possible solutions in the decentralized setting

If the states have incentives to under-provide health care, the federal government might choose a point in the South West quadrant, as illustrated in Figure 1. This case illustrates a situation where the federal government has induced the states to increase their expenditure on health care, compared to point 2, by reducing the sickness benefit. If the low state tax rates, associated with low expenditure on health care, result in sufficiently low social cost for financing the sickness benefit and the slope of the states' reaction functions is sufficiently flat, the federal government will choose a point in the South East quadrant. Low state tax rates could reduce the social cost of financing the sickness benefit by reducing the share of mimickers and the workers' marginal benefit of income, and by increasing the number of hours each worker supplies. However, low expenditure on health care implies a high share of the population being absent from work due to illness, which increases the social cost of financing the sickness benefit. If this effect is strong enough and the slope of the states' reaction functions is steep enough, the federal government might end up choosing a point in the North West quadrant, which means that e will be above e^*, despite the states having incentives to under-provide health care for any given level of B.

To conclude, given the general form of the model we can not determine how the level of health care provided in the decentralized setting will be in relation to the second best level, except when the externalities facing the state governments cancel out exactly. Whether health care will be under- or over-provided in the other situations depends not only on the sign of the fiscal externality facing the states, but also on the social costs of financing the sickness benefit, which affects the form of the federal government's indifference curves, and on the slope of the states' reaction functions.

3.3 Implementing the second best solution

This section examines different ways that the unitary nation optimum can be achieved by changing the governments' responsibilities or policy instruments. A self-evident solution is to abolish the federal structure, either by transferring the responsibility of providing health care to the federal level or by transferring the responsibility of providing the sickness benefit to the state level. In this model these two solutions will be equally effective, but in reality both solutions might be unfeasible for constitutional, political or other reasons not accounted for in the model.

It is also possible to implement the second best solution while retaining the federal structure. If the federal government is given the possibility of deciding a positive or negative intergovernmental transfer, R, the state and the federal governments' budget constraints become

$$(1 - a(e))(1 - \hat{a})twl - e + R = 0 \tag{16}$$

and

$$(1 - a(e))(1 - \hat{a})Twl - (a(e) + \hat{a}(1 - a(e)))B - R = 0,$$

respectively. The federal government now has three policy instruments at its disposal, T, B and R. It can observe the true marginal cost of public funds and the preference of the individuals and is therefore able to set $B = B^*$. The federal government is also able to observe the optimal value of τ and to design the intergovernmental transfer so that the state governments will internalize the effects that their decisions have on the federal government. This is described in Proposition 2.

Proposition 2 *If the federal government remits all its tax revenues back to the states, but makes each state transfer back an amount equal to the expenditure on the optimal sickness benefit in that state, then $t = \tau^* - T$ and $e = e^*$ will solve the state governments' optimization problem. The intergovernmental transfer can be written as*

$$R = (1 - a(e))(1 - \hat{a})(Twl + B^*) - B^*, \tag{17}$$

where $(1 - a(e))(1 - \hat{a})$ is the share of workers, Twl is the federal tax income per worker and B^* is the optimal sickness benefit.[15]

Proof. By substituting equation (17) into equation (16) the states optimization problem becomes

$$\max_{t,e}\ (1-a(e))(1-\hat{a})\ v(w(1-t-T), h^h) + a(e)\widetilde{v}(B, h^s) + (1-a(e)) \int_{i=0}^{\hat{a}} [\widetilde{v}(B, h^h) - g(m_i)]di$$

s.t.

$$(1 - a(e))(1 - \hat{a})twl - e + (1 - a(e))(1 - \hat{a})(Twl + B^*) - B^* = 0,$$

where $a(e)$, $\hat{a}$ and l are defined as before. The Lagrangian is written

$$\begin{aligned} L \quad = \quad & (1 - a(e))(1 - \hat{a})\ v(w(1 - t - T), h^h) + a(e)\widetilde{v}(B, h^s) \\ & +(1 - a(e)) \int_{i=0}^{\hat{a}} [\widetilde{v}(B, h^h) - g(m_i)]di \\ & +\gamma^s\{(1 - a(e))(1 - \hat{a})(t + T)wl - (a(e) + \hat{a}(1 - a(e)))B^* - e\}. \end{aligned}$$

Since $\tau = t + T$ by definition, this Lagrangian is identical with that for the social optimization problem in subsection 2.1.

The fact that T can be set at any level is a result of t and T entering additively in the individuals' utility functions. The choice of T will not affect the

[15] Aronsson and Wikström (2003) derived a similar result, in the context of risk-sharing in an economic federation.

policy rule, as such, but the size of the intergovernmental transfer. That is, T is a superfluous policy instrument. However, if the tax rates are constrained to be non-negative, T must be $0 \leq T \leq \tau^*$. To understand why the transfer will make the externalities facing the states vanish, notice that equation (17) is identical with the federal government's budget constraint without an intergovernmental transfer, equation (7), when $B = B^*$. Therefore, the states will take the same decisions as they would have taken if they directly took the effect of their decisions on the federal government's budget constraint into account.

A special case is when $T = 0$, which allows the transfer to be written

$$R = -(a(e) + \hat{a}(1 - a(e)))B^*. \tag{18}$$

The transfer will in this case be negative and equal to the expenditure on the optimal sickness benefit in each state. The size of the negative transfer depends on each state's decisions regarding e and t and not only has the objective to finance the federal government's expenditure on the sickness benefit, but also to correct the states' incentives. This result differs from that presented by for example Boadway and Keen (1996). They claim that if all rents and tax powers are allocated to the states, the vertical externality will vanish and the sole objective of the transfer will be to finance the federal expenditure. The different result is caused here by the fact that the federal government has at its disposal a redistributive policy instrument, B, which was not present in Boadway and Keen (1996), and the fact that the need for redistribution is directly affected by the decisions taken by the states.

Moral hazard among individuals (captured by the parameter $\hat{a}$) does not affect the different ways the second best solution can be implemented, but does affect the size of the optimal transfers. This can be seen in equations (17) and (18).[16]

4 Discussion

This paper addresses provision of health care and redistribution, in terms of a sickness benefit, in an economic federation. The analysis is based on a model

[16] Similarly, Proposition 2 holds even if mimickers' utilities are assigned a lower social weight, but the size of the optimal transfers would be affected since this would affect the values of the parameters.

where the state governments provide health care and the federal government provides a sickness benefit and both levels of government tax labor income.

The results show that the states can either have an incentive to under- or over-provide health care. The federal government can induce the states to increase (reduce) their expenditure on health care, by reducing (increasing) the sickness benefit and the federal tax rate. The results also demonstrate that the federal government can induce the state governments to internalize the effects that their decisions will have on the federal government's budget constraint. This can be done by introducing an intergovernmental transfer, where the federal government remits all its tax revenues to the states, but makes each state transfer back an amount equal to the expenditure for the optimal sickness benefit in that state. In this model, the vertical fiscal externality will not vanish even if all powers of taxation are assigned to the states. This result differs from previous ones presented in the literature and is caused by the fact that the states' decisions in this model directly affect the federal government's expenditure.

The results from this paper can be generalized to other state financed programs that reduce the number who receive federal transfers, for example labor market programs, and may inform policy makers on how to reduce the misallocation of resources associated with such programs.

One important assumption in the paper is that no horizontal externalities exist. In reality, labor mobility will give rise to horizontal externalities. If labor is mobile, expenditure on health care in a state may attract sick individuals to that state and discourage workers, due to the higher tax rate necessary for financing the increased expenditure. Other states will therefore benefit from a state's increased expenditure, which implies that labor mobility gives rise to a positive horizontal externality.[17] Including these positive horizontal externalities in the analysis would therefore increase the probability of health care being under-provided in the decentralized setting, but it will not change the way the federal government can influence the states' decisions and the general result that the unitary nation optimum can be implemented through an intergovernmental transfer.

[17] The possibility of patients seeking treatment in another state than that of resident, will also cause a horizontal externality, if the state of resident does not have to fully reimburse the state which treated the patient.

Appendix

Define

$$F = [1 + (1 - \hat{a})twla'(e)],$$

$$G = (1 - a(e))\left[(1 - \hat{a})(wl - tw^2 l') - twl\frac{\partial \hat{a}}{\partial t}\right]$$

and note, by comparing with equation (9), that $\frac{F}{G} = \frac{dt}{de}$. Using these definitions, the derivative of equation (9) with respect to T can be written

$$\begin{aligned}\frac{\partial(dt/de)}{\partial T} &= G\left[-\frac{\partial \hat{a}}{\partial T}twla'(e) - (1-\hat{a})a'(e)tw^2 l'\right] \\ &\quad -F\{(1-a(e))[(1-\hat{a})(-w^2 l' + tw^3 l'') \\ &\quad -\frac{\partial \hat{a}}{\partial T}(wl - tw^2 l') - twl\frac{\partial^2 \hat{a}}{\partial t \partial T} + tw^2 l'\frac{\partial \hat{a}}{\partial t}]\}/G^2. \qquad (19)\end{aligned}$$

By substituting equation (19) into equation (14) and rearranging, the latter equation can be written

$$\frac{de}{dT} = -\frac{\begin{array}{c} a'(e)(1-\hat{a})v'w + (1-a(e))(1-\hat{a})v''w^2\frac{dt}{de} \\ +(1-a(e))v'w/G\frac{\partial \hat{a}}{\partial T}[F + (1-\hat{a})twla'(e) \\ -F\frac{(1-\hat{a})(1-a(e))(wl - tw^2 l')}{G}] \\ +(1-a(e))(1-\hat{a})v'w\{(1-\hat{a})a'(e)tw^2 l'/G \\ +F(1-a(e))\left[(1-\hat{a})(-w^2 l' + tw^3 l'') - twl\frac{\partial^2 \hat{a}}{\partial t \partial T} + tw^2 l'\frac{\partial \hat{a}}{\partial t}\right]/G^2\} \end{array}}{\delta}. \qquad (20)$$

The terms in the four last rows in the numerator of equation (20) describe how the state's perception of the marginal cost of public funds is affected by the federal tax rate, through its effect on the share of workers in the population and its effect on the relationship between t and e. As will be demonstrated, these four rows are jointly negative given the assumptions made. G and F are positive according to the state's first order condition. The quotient in the third row is larger than one, which guarantees that the second and third rows are jointly negative. In the fifth row, the term $tw^2 l'\frac{\partial \hat{a}}{\partial t}$ is dominated by the term $-(1-\hat{a})w^2 l'$, since the assumption that the state's tax revenue is strictly increasing in its tax rate, requires that $t\frac{\partial \hat{a}}{\partial t} < (1-\hat{a})$. This guarantees also that the fourth and fifth rows are jointly negative.

The derivative of equation (9) with respect to B is written

$$\begin{aligned}\frac{\partial(dt/de)}{\partial B} &= \{-G\frac{\partial \hat{a}}{\partial B}twla'(e)\\ &\quad +F(1-a(e))\left[\frac{\partial \hat{a}}{\partial B}(wl-tw^2l')+twl\frac{\partial^2 \hat{a}}{\partial t\partial B}\right]\}/G^2. \qquad (21)\end{aligned}$$

By substituting equation (21) into equation (15) and rearranging, the latter equation can be written

$$\frac{de}{dB} = -\frac{\begin{matrix} a'(e)\left[\frac{\partial \widetilde{v}(B,h^s)}{\partial B}-\hat{a}\frac{\partial \widetilde{v}(B,h^h)}{\partial B}\right] \\ +(1-a(e))v'w/G\frac{\partial \hat{a}}{\partial B}[F+(1-\hat{a})twla'(e) \\ -F\frac{(1-a(e))(1-\hat{a})(wl-tw^2l')}{G}] \\ -(1-a(e))(1-\hat{a})v'wF(1-a(e))twl\frac{\partial^2 \hat{a}}{\partial t\partial B}/G^2 \end{matrix}}{\delta}. \qquad (22)$$

The terms in the last three rows in the numerator of equation (22) describe how the state's perception of the marginal costs of public funds is affected by the sickness benefit. Given the assumptions made, these terms are jointly negative for the same reason as described above.

$\frac{dt}{dT}$ and $\frac{dt}{dB}$ can be derived in a similar manner. Maximizing the states' objective functions with respect to t, letting the state's expenditure on health care be defined subsequently by its budget constraint, gives the first order condition

$$-\left[(1-\hat{a})V^h-V^s+\int_{i=0}^{\hat{a}}\widehat{V_i^h}di\right]a'(e)\frac{de}{dt}-(1-a(e))(1-\hat{a})v'(w(1-t-T),h^h)w=0. \qquad (23)$$

Differentiating equation (23) with respect to T and t, holding B fixed gives

$$\frac{dt}{dT} = -\frac{\begin{matrix} a'(e)(1-\hat{a})v'w\frac{de}{dt}-\left[(1-\hat{a})V^h-V^s+\int_{i=0}^{\hat{a}}\widehat{V_i^h}di\right]a'(e)\frac{\partial(de/dt)}{\partial T} \\ +(1-a(e))(1-\hat{a})v''w^2+(1-a(e))v'w\frac{\partial \hat{a}}{\partial T} \end{matrix}}{\zeta}. \qquad (24)$$

ζ is the differential of equation (23) with respect to t and is therefore negative, given that the state's objective function is concave in t. The first term in the numerator of equation (24) shows that an increased federal tax rate will reduce the workers utility, reducing the state's incentive to raise taxes to finance the treatment of sick individuals. The second term demonstrates that this incentive is further reduced by the fact that an increased federal tax rate reduces the total number of hours worked, reducing the amount of health care that can be financed by a given state tax. Further more, an increased federal tax rate will increase the workers marginal benefit of income, increasing the social cost of raising the taxes. The fourth term shows that the social cost of raising taxes is reduced by the fact that an increased federal tax rate reduces the number of workers that are affected by the tax rate. By substituting equation (25)

$$\begin{aligned}\frac{\partial(de/dt)}{\partial T} &= \{F(1-a(e))[(1-\hat{a})(-w^2l' + tw^3l'') \\ &\quad -\frac{\partial\hat{a}}{\partial T}(wl - tw^2l') - twl\frac{\partial^2\hat{a}}{\partial t\partial T} + tw^2l'\frac{\partial\hat{a}}{\partial t}] \\ &\quad +G\frac{\partial\hat{a}}{\partial T}twla'(e) + G(1-\hat{a})tw^2l'a'(e)\}/F^2 \qquad (25)\end{aligned}$$

into equation (24) and rearranging using equation (23), it can be seen below that the fourth term in equation is dominated by the second one.

$$\frac{dt}{dT} = -\frac{\begin{array}{c} a'(e)(1-\hat{a})v'w\frac{de}{dt} + (1-a(e))(1-\hat{a})v''w^2 \\ -a'(e)\left[(1-\hat{a})V^h - V^s + \int\limits_{i=0}^{\hat{a}} \widehat{V_i^h}di\right] \\ \{F(1-a(e))\left[(1-\hat{a})(-w^2l' + tw^3l'') - twl\frac{\partial^2\hat{a}}{\partial t\partial T} + tw^2l'\frac{\partial\hat{a}}{\partial t}\right] \\ +G(1-\hat{a})tw^2l'a'(e)\}/F^2 \\ +(1-a(e))v'w\frac{\partial\hat{a}}{\partial T}/F\left[F + (1-\hat{a})twla'(e) - F\frac{(1-a(e))(1-\hat{a})(wl-tw^2l')}{G}\right]\end{array}}{\zeta}$$

Differentiating equation (23) with respect to B and t, holding T fixed gives

$$\frac{dt}{dB} = -\frac{\begin{array}{c} a'(e)\left[\frac{\partial\tilde{v}(B,h^s)}{\partial B} - \hat{a}\frac{\partial\tilde{v}(B,h^h)}{\partial B}\right]\frac{de}{dt} \\ -a'(e)\left[(1-\hat{a})V^h - V^s + \int\limits_{i=0}^{\hat{a}} \widehat{V_i^h}di\right]\frac{\partial(de/dt)}{\partial B} \\ +(1-a(e))v'w\frac{\partial\hat{a}}{\partial B}\end{array}}{\zeta} \qquad (26)$$

The first term in the numerator of equation (26) shows that an increased sickness benefit will increase the utility of people on sick leave. Given the assumption stated in Section 3.1, this will reduce the state's incentive to raise taxes to finance health care. The second term demonstrates that this incentive is further reduced by the fact that a higher sickness benefit reduces the total number of hours worked, reducing the amount of health care that can be financed by a given state tax rate. The last term shows that the social cost of raising taxes is reduced by the fact that an increased sickness benefit reduces the number of workers that are affected by the tax rate. By substituting equation (27)

$$\begin{aligned} \frac{\partial(de/dt)}{\partial B} &= \{-F(1-a(e))\left[\frac{\partial \hat{a}}{\partial B}(wl - tw^2 l') + twl\frac{\partial^2 \hat{a}}{\partial t \partial B}\right] \\ &\quad +G\frac{\partial \hat{a}}{\partial B} twla'(e)\}/F^2 \end{aligned} \tag{27}$$

into equation (26) and rearranging using equation (23), it can be seen below that the last term in equation (26) is dominated by the second one.

$$\frac{dt}{dB} = -\frac{\begin{array}{c} a'(e)\left[\frac{\partial \widetilde{v}(B,h^s)}{\partial B} - \hat{a}\frac{\partial \widetilde{v}(B,h^h)}{\partial B}\right]\frac{de}{dt} \\ +a'(e)\left[(1-\hat{a})V^h - V^s + \int\limits_{i=0}^{\hat{a}} \widehat{V_i^h} di\right] \\ (1-a(e))twl\frac{\partial^2 \hat{a}}{\partial t \partial B}/F \\ +(1-a(e))v'w/F\frac{\partial \hat{a}}{\partial B}\left[F + (1-\hat{a})twla'(e) - F\frac{(1-a(e))(1-\hat{a})(wl - tw^2 l')}{G}\right] \end{array}}{\zeta}$$

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The Effect of Health Care Expenditure on Sickness Absence

David Granlund*
Department of Economics
Umeå University
SE-901 87 Umeå
Sweden

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Abstract

Based on data from a panel of the Swedish municipalities during 1993-2004, the effects of public health care expenditure on absence from work due to sickness or disability were studied using an instrumental variable method. Public health care expenditure had no significant effect on absence due to sickness or disability and the standard errors were small enough to rule out all but a minimal effect. The same result was obtained when separate estimates were done for men and women and for absence due to sickness and disability.

Key words: health care expenditure; sick leave; disability; worker absenteeism; dynamic panel data models; endogeneity

JEL classification: H51; I12; J22

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1 Introduction

Rising health care costs in most industrialized countries have increased the importance of evaluating the effects of health care expenditure. In the past few decades, assessing its effects on health outcomes has become a central question in the context of health care cost containment in most developed countries (Nixon and Ulmann, 2006). From a Swedish perspective it is especially interesting to estimate the relationship between health care expenditure and absenteeism, since according to OECD (2005) "Sweden's single biggest economic problem is the high number of people absent from work due to sickness or disability".[1]

The effects of health care programs on the absence of specific patient groups have been studied previously.[2] However, to my knowledge, the effect of aggregated health care expenditure on absence has never been studied. The purpose of the present study was to estimate how aggregated public expenditure on health care affects absence from work due to sickness or disability.

The literature on absence includes estimations of the effects of individuals' health status: Paringer (1983) found perceived health status to be an important predictor of hours lost from work, which was supported by Primoff Vistnes (1997), who also reported statistically significant effects of obesity and smoking on the likelihood for women's absence. The literature also provides massive support for that economic incentives affect absence, for example the following four studies using Swedish data; Johansson and Brännäs (1998), Johansson and Palme (2002, 2005); Henrekson and Persson (2004).[3]

Whether increased aggregated expenditure on health care or increased access to health care actually improves the health status of the population in industrialized countries is still an open question. Higher expenditure on health care could lead to better health among the population by reducing waiting times for medical care or improving procedures. But according to Nixon and Ulmann's (2006) review of the literature on health care expenditures and health outcomes, cross-country studies have found limited or no relationship between health care expenditure and mortality rates. On the other hand, Crémieux et al. (1999)

[1] According to Statistics Sweden and The Swedish Social Insurance Agency, 5.2 percent of employee working hours in Sweden were lost due to sickness absence in 2004; at the same time 8.1 percent of the population aged 16 to 64 were on disability pension.

[2] Absence will be used throughout this paper to mean absence from work due to either sickness or disability.

[3] Brown and Sessions (1996) review the literature on absence more broadly.

found that higher health care expenditure among Canadian provinces reduced male and female infant mortality and increased life expectancy. They explained the different results by the inherent heterogeneity associated with cross-country studies. Lichtenberg (2004) analyzed time-series of life expectancy in the United States and found that both public health care expenditure and research and development expenditure on pharmaceuticals had positive effects. Aakvik and Holmås (2006) found no effect of the total number of general practitioners per capita on mortality rates in Norwegian municipalities, but found a negative effect of the number of contracted general practitioners. Brook et al. (1983) reported on the Rand Health Insurance Experiment, a controlled trial in the United States where families were randomly assigned insurance plans. One group received all their medical care free of charge and, as a consequence, used more than the other groups. Despite this, the only statistically significant effects were improvements in health for those with poor vision and for low-income persons with high blood pressure. However, the study included only people aged 14 to 61 who were free of disability that precluded work.

Except for the rare occasions when a randomized controlled trial is preformed, determining the effect of health care expenditure and access to health care is complicated by the context in which decisions regarding health care are taken. Health care expenditure is partly determined by the perceived need for it, which in turn may be affected by absence. Therefore an instrumental variable estimator was used in this study in an attempt to determine the causal link between health care expenditure and absence.

Most previous studies have evaluated the effect of health care expenditure on mortality rates and life expectancy. Therefore, this paper contributes to the literature by examining the effect on absence due to sickness or disability, which can be expected to be correlated with individuals' health related quality of life. The analysis was based on municipality-level data from Sweden. As demonstrated by Granlund (2007), the results may also help determining the sign of a vertical fiscal externality that arises when a lower level of government provides health care, the central government provides a sickness benefit and both levels tax labor income. Health care is more likely to be over-provided by the local governments the smaller effect health care expenditure has on absence. The article also explains why the lower level of government has a weak incentive to reduce absence. In practice, this may result in a relatively small share of total health care expenditure being focused on reducing absence in a country like Sweden. The results from research in this field may also inform policy makers in their decision regarding the level of expenditure on health care. The main finding in this paper is that health care expenditure had no statistically significant effect on absence and that, under all circumstances, the possible effect was small.

The next section outlines the theory, while section 3 presents the empirical analysis. The data are discussed in section 3.1 and the empirical specification in section 3.2, while section 3.3 contains the results. Finally, in section 4 the paper's conclusions are presented.

2 Theoretical outline

To later be able to specify the empirical model, we need an absence function for municipalities which includes the health status of the population as well as a health production function. To derive the first of these, we first have to analyze what determines whether individuals will prefer to be absent from work.

An individual's instantaneous utility can be expressed as $u_t = u(c_t) - f(h_t, j_t) - g(s_{i,t-1}, s_{-i,t-1})$, where c_t is consumption at time t; effort (f) depends on health status (h_t) and work conditions (j_t); and the disutility of absence $(-g)$ depends on one's own prior period absence $(s_{i,t-1})$ and that of others $(s_{-i,t-1})$. It is assumed that effort decreases with good health and work conditions, while the disutility of absence decreases with both one's own and others' prior period absence. This represents internal habit formation, that one might be more likely to be absent this period for a given health and other variables if one was absent the last period, and external habit formation, that high previous absence of others in one's surroundings might reduce the social cost of current absence. Naturally, for absent individuals the effort is zero and for those working the disutility of absence is zero.

For simplicity let's assume that individuals have no access to capital markets and only choose whether to work or be absent. When an individual works $c_t = w_t(1 - \tau_t)$, where w is labor income and τ is the tax rate on labor income. For individuals absent from work $c_t = B_t(1 - \tau_t)$, where B is a sickness benefit. Assuming myopic behavior, i.e. that individuals neglect the effect of current absence on their future utility, individual's will prefer to be absent if

$$u(B_t(1-\tau_t)) - g(s_{i,t-1}, s_{-i,t-1}) > u(w_t(1-\tau_t)) - f(h_t, j_t). \qquad (1)$$

Medical doctors and sickness insurance officials are usually involved in deciding whether a sickness benefit will be allowed. According to the Swedish regulations during the study-period, individuals were entitled to sickness benefits if their capacity to work was sufficiently reduced due to poor health and a doctor's certificate was usually required for sick leave extending one week.[4] For long-term sick leave, the capacity to perform other assignments would also be taken into consideration.[5] Nevertheless, also the variables B, τ, $s_{i,t-1}$, $s_{-i,t-1}$ and w can be expected to affect absence. First, since doctors and insurance officials cannot observe ability to work perfectly, they must rely in part on information provided by the individual, which may be affected by the individual's incentives. Second, doctors may also consider what their patients prefer, and while insurance officials can deny sickness benefits to individuals with a doctor's certificate, they mostly follow the doctor's recommendation.[6] The absence function for municipalities can therefore be written

$$\mathbf{s_t} = s(\mathbf{B_t}(1-\boldsymbol{\tau}_\mathbf{t}), \mathbf{w_t}(1-\boldsymbol{\tau}_\mathbf{t}), \mathbf{s_{t-1}}, \mathbf{j_t}, \mathbf{h_t}), \qquad (2)$$

where $\mathbf{B}, \boldsymbol{\tau}, \mathbf{w}, \mathbf{s}, \mathbf{j}, \mathbf{h}$ are in bold to indicate that they describe the situation for all individuals of working age in a municipality.

The aggregated health status in a municipality at period t can be written

$$\mathbf{h_t} = h(\mathbf{h_{t-1}}, \mathbf{e_t}, \mathbf{r_t}, \mathbf{X_t}), \qquad (3)$$

where $\mathbf{e}$ is public health care expenditure per capita; $\mathbf{r}$ represents the health risks that individuals are exposed to, for example at their workplaces; and $\mathbf{X}$ denotes demographic characteristics, like age and gender.

[4] Since October 1, 1995, a doctor's certificate has been required to receive benefits after the seventh calendar day of absence. Before that, the Swedish Social Insurance Agency could require a doctor's certificate for sick leave lasting over four weeks, and also earlier in some cases (Proposition 1994/95:147; Lag 1962:381).

[5] In certain circumstances, also the individual's age and education etc. were allowed to influence the judgment regarding capacity to work (Lag 1962:381).

[6] Blomqvist (1991) and Shortell (1998) discussed physicians multiple accountabilities. Based on a survey of 4,200 physicians active in Sweden (response rate 58 percent), Arrelöv (2006) similarly reports that 65 percent of physicians consider the patient's motivation for returning to work when assessing the extent of allowable sick leave.

3 Empirical analysis

3.1 Data description

The present study was based on yearly municipality-level data on absence for the period 1993 to 2004. There were 286 municipalities in 1993, yielding 3,432 observations.[7] Data on public health care expenditure are primarily available at county level, where responsibility for health care provision lies. There were 23 counties in 1993, shrinking to 20 in 2004. In addition there were three municipalities (Malmö, Göteborg and Gotland) which did not belong to any county but provided health care themselves in 1993. By 2004 only Gotland remained in this category.

Table 1 (below) gives descriptive statistics of the variables used in this study while Table A1 in Appendix A defines them and gives data sources. Figure 1 and 2 (also in Appendix A) provide box plots for two of the most central variables in this study. The first six variables describe absence from work and cover all employees and self-employed in Sweden, since they were all automatically insured in the social insurance system. *Sickness* is the average number of days of absence from work due to sickness during a year for insured individuals in the ages of 16 to 64. *Disability* and *Rehab* are the corresponding numbers of days on disability/early retirement pension and days of absence due to rehabilitation, respectively, and absence, s, is the sum of these three variables. The original absence data lacked information for some observations and did not include days compensated by employers, which over the study-period changed between the first 14, 21, and 28 days of each absence spell. In Appendix B it is described how data from other sources were used to adjust the absence variables to always correspond to absence from the 15th day of each spell, as well as how missing data in these variables were handled.

The most common reasons for both sickness and disability absence were illness in locomotion organs and mental illness. Those whose capacity to work was expected to be sufficiently reduced for a long time could receive disability/early retirement pension, usually preceded by a long period of sick leave (Riksförsäkringsverket, 2004; Lag 1962:381). During the study-period, the compensation levels in the social insurance system ranged from 75 to 90 percent of

[7] By 2004 there were 290 municipalities, but data from the new municipalities were aggregated according to 1993 boundaries.

the income from the second day of absence, but with a cap at a certain level of income. At first less than 10 percent of the insured were affected by this cap, but by 2004 22 percent were (Henrekson and Persson, 2004; the Swedish Social Insurance Agency).

Public health care expenditure (e) was defined as each county's per capita operating costs on health care, excluding expenditure on dental care and pharmaceuticals. Of this roughly 2 percent was patients' co-payments for public health care. During the study-period total health care expenditure constituted 7.5-8.5 percent of Sweden's GDP, of which 11-15 percent was for pharmaceuticals and 8-10 percent for dental care. Pharmaceuticals were excluded from the study because they were paid by the central government until 1998, and dental care was excluded since it might have a quite different effect on absence compared to other health care services. Public expenditure accounted for approximately 95 percent of the total non-dental, non-pharmaceutical health care expenditure in Sweden (Socialstyrelsen, 2006). The variable (e) of course includes expenditure on the entire population (not just those of working age), but adjustments for variations in county age-composition were made using micro-data of health care consumption. Appendix C describes how this was done in order to create a variable describing age-adjusted per capita public health care expenditure, denoted $eadj$.

w is the average labor income of the non-absent population of working age (16 to 64 years of age) in each municipality. τ^M and τ^C are the proportional municipality and county income tax rates, respectively, and τ^{MC} is the sum of them.

The variables $Women$ to $Pop6064$ describe the shares in the population of working age which belong to each demographic group. $El.School$, $HighSchool$ and $University$ denote the shares of the working age population with different educational levels, described in Table A1 in Appendix A. $SocM$ and $SocC$ denote the fraction of each municipality and county parliament, respectively, represented by socialist parties. Finally, $PolmajM$ and $PolmajC$ are dummy variables which take the value 1 if either one of the two traditional Swedish political blocks has own majority in the municipality and county parliament, respectively.

Table 1. Descriptive statistics

	1994**		2004		1993-2004	
Variable	Mean	Std.dv.	Mean	Std.dv.	Mean	Std.dv.
$Sickness$	10.46	1.84	15.55	3.30	13.44	4.08
$Disability$	29.08	8.06	31.26	7.20	28.72	7.40
$Rehab$	1.18	0.49	1.01	0.46	0.89	0.45
s	40.72	9.23	47.82	9.14	43.05	9.55
s_{women}	45.97	10.06	58.39	11.00	51.18	11.50
s_{men}	35.97	9.02	37.88	8.20	36.03	8.44
e^*	10.84	1.25	15.76	0.92	12.86	2.11
$eadj^*$	10.68	1.02	15.56	0.81	12.72	2.04
w	159.70	18.25	215.62	23.65	186.39	29.58
τ^m	19.24	1.80	21.45	1.26	20.70	1.72
τ^{c*}	11.09	1.34	10.43	0.68	10.24	1.11
τ^{mc}	30.34	1.12	31.84	0.91	30.94	1.21
$Women$	0.49	0.01	0.49	0.01	0.49	0.01
$Pop1639$	0.50	0.02	0.45	0.04	0.47	0.04
$Pop4049$	0.24	0.01	0.21	0.01	0.23	0.02
$Pop5054$	0.10	0.01	0.11	0.01	0.12	0.01
$Pop5559$	0.08	0.01	0.13	0.01	0.10	0.02
$Pop6064$	0.08	0.01	0.10	0.01	0.09	0.01
$El.School$	0.33	0.06	0.25	0.04	0.30	0.06
$HighSchool$	0.59	0.03	0.63	0.03	0.61	0.04
$University$	0.07	0.04	0.12	0.05	0.09	0.05
$SocM$	0.51	0.12	0.47	0.11	0.48	0.12
$SocC^*$	0.45	0.07	0.49	0.04	0.50	0.06
$PolmajM$	0.73	0.47	0.71	0.45	0.72	0.45
$PolmajC^*$	0.93	0.24	0.44	0.50	0.66	0.48

*Indicates that the variable is measured at county-level instead of municipality-level.
** Descriptive statistics are reported for 1994 since data on s^{women} and s^{men} were not available for 1993.

3.2 Empirical specification

The empirical specification of the municipal absence function, i.e. equation (2), can be written

$$\begin{aligned} s_{it} &= \boldsymbol{\beta}_1 w_{it}(1-\tau_{it}^{mc}) + \beta_2 s_{i,t-1} \\ &\quad + \sum_{l=1}^{2} \eta_l Edu_{lit} + \boldsymbol{\beta}_3 h_{it} + y_t + \mu_i + \varepsilon_{it}. \end{aligned} \tag{4}$$

$w_{it}(1-\tau_{it}^{mc})$, the net labor income in municipality i at time t, was included to capture the monetary incentive of remaining at work for the marginal worker. In equation (2) absence was also affected by the sickness benefit net of taxes, but this was left out from the specification since sickness benefits are a function of labor income and since observed values of it to a higher degree than the observed values of w depend on the composition of those on absence. Hence, a relatively high fraction of the variation in B does not correspond with variation in the monetary incentives to remain at work for the marginal worker.

The educational variables, Edu_l, $(l = 1, 2)$, were used as proxies for work conditions. These variables may also capture the effect caused by that employment contracts differ among different type of jobs in respect to stipulated number of hours and flexible hours. Effects of contracts have been highlighted as major explanations to absence in previous economic literature (see e.g. Brown and Sessions, 1996). Year-specific fixed effects (y_t) were included to capture “national variables” such as business cycle effects on absence. Municipality-specific fixed effects (μ_i) were included to capture time invariant heterogeneity among the municipalities which might be correlated with the regressors. The other two variables, s_{t-1} and h, were motivated in the theoretical outline.

The empirical specification of the municipal health production function, i.e. equation (3), can be written

$$\begin{aligned} h_{it} &= h_{i,t-1} + \gamma_2 eadj_{it} + \sum_{l=1}^{2} \delta_l Edu_{lit} + \sum_{n=1}^{4} \zeta_n \Delta Pop_{nit} \\ &\quad + \sum_{n=1}^{4} \kappa_n Pop_{nit} + \gamma_3 Women_{it} - \delta_t, \end{aligned} \tag{5}$$

where h_{t-1} denotes lagged health status and $eadj$ is age-adjusted health care expenditure. Since lagged health status was included as an explanatory variable, the purpose of the other explanatory variables was to capture changes in the health status, rather than the level of it. As such, Edu_l, $(l = 1, 2)$ were used as proxies for the health risks that people in the municipalities are exposed to during the year at their work place. $\Delta Pop_{nt} = Pop_{nt} - Pop_{n,t-1}$, $(n = 1, 2...4)$ describe the change in the age-composition of the population in working age, whereas Pop_n, $(n = 1, 2...4)$ and $Women$ describe the demographic composition of the population in working age. The demographic variables might enter the equation in differences since demographic groups might differ in health status and that changes in these variables therefore lead to changes in the population's health status. Demographic variables might enter in levels since demographic groups might have different development of their health status over time.[8] δ denotes depreciation of the health status and was allowed to vary over time but not over municipalities. Not allowing the depreciation of the health status to vary over municipalities in other ways than that captured by the demographic and the educational variables was of course a restriction. This restriction was imposed since health status is hard to measure and that it therefore is difficult to estimate how the depreciation depends on the level of this variable.

Differentiating equation (5) and substituting it into a differentiated version of equation (4) yields

$$\begin{aligned}\Delta s_{it} \quad = \quad & \boldsymbol{\beta}_1 \Delta(w_{it}(1-\tau_{it}^{mc})) + \boldsymbol{\beta}_2 \Delta s_{i,t-1} \sum_{l=1}^{2} \eta_l \Delta Edu_{lit} + \\ & + \boldsymbol{\beta}_3 \{\gamma_2 eadj_{it} + \sum_{l=1}^{2} \delta_l Edu_{lit} + \sum_{n=1}^{4} \zeta_n \Delta Pop_{nit} \\ & + \sum_{n=1}^{4} \kappa_n Pop_{nit} + \gamma_3 Women_{it}\} + Y_t + \Delta\varepsilon_{it}, \qquad (6)\end{aligned}$$

where $Y_t = y_t - y_{t-1} - \boldsymbol{\beta}_3 \delta_t$. This equation, on which all empirical specifications will be based, shows how the effect of health care expenditure on absence can be estimated without having to include proxies for health status. The new error term, $\Delta\varepsilon_t = \varepsilon_t - \varepsilon_{t-1}$, is by construction correlated with the lagged dependent variable, $\Delta s_{t-1} = s_{t-1} - s_{t-2}$, which is then endogenous. Also expenditure on health care might be endogenous since, ceteris paribus, an increase in absence might cause the counties to increase it. Beyond that, a negative health

[8] $\Delta WomenWA$ was left out since it is has no statistically significant effect on absence.

shock, not captured by any of the other explanatory variables of equation (6), might cause an increase in both absence and health care expenditure. However, the endogeneity problem is probably reduced since the dependent variable was measured at the municipal level, whereas health care expenditure was decided at the county level. Since county tax revenue was used to finance health care the tax rate might also be endogenous, and if the labor income of the marginal non-absent individual differs from the average labor income of the non-absent population, w will be endogenous as well.

The endogeneity problem was addressed by instrumenting Δs_{t-1}, $eadj$ and $\Delta(w(1-\tau^{mc}))$ with the closest lags uncorrelated with the new error term, namely Δs_{t-2}, $eadj_{t-1}$, and $\Delta(w_{t-2}(1-\tau^{mc}_{t-2}))$; and $\Delta PolmajM$, $\Delta PolmajC$, $\Delta SocM$, and $\Delta SocC$ were included as additional instruments.[9,10] These four last instruments are expected to correlate with the tax and expenditure decisions of the municipal and county governments and were primarily included to avoid problems with weak instruments for $\Delta(w(1-\tau^{mc}))$; but $\Delta PolmajC$ and $\Delta SocC$ could also be expected to strengthen the instrument set for health care expenditure.[11]

Based on the theoretical outline it is reasonable to expect $\Delta(w(1-\tau^{mc}))$ to have a negative impact on absence. Descriptive statistics from Sweden states that younger and better educated individuals were less likely to be absent, which formed my expectation about the coefficients for the differenced educational and demographic variables. During the period under study, absence has increased much more for women than for men, more for individuals aged 50 to 59 than for others and the absence has decrease for individuals aged 60 to 64. This formed my expectations for the coefficient for $Women$ and the three highest age-groups in levels, whereas I had no prior expectation regarding the remaining age-group and Edu_l , $(l = 1, 2)$. Previous absence was expected to have a positive influence on current absence, as explained in the theoretical outline. Lastly, public health care expenditure was anticipated to have negative or no impact on absence. As

[9] These instruments will be uncorrelated with the new error term if ε_t and ε_{t-1} are uncorrelated, which is likely since the lagged dependent variable is included in the model. ε_t and ε_{t-1} are correlated if $\Delta\varepsilon_t$ and $\Delta\varepsilon_{t-2}$ are correlated, which can be tested.

[10] In the static specifications, Δs_{t-2} was not included.

[11] Similar variables were used by Aronsson et al. (2000) in a regression of municipal tax base, but they used a Herfindal-Index of political fragmentation instead of $\Delta PolmajM$ and $\Delta PolmajC$ as here.

reported in the introduction, some previous research has reported no or very limited effects of aggregated health care expenditure on the health status of the population. Moral hazard problems is one set of explanations to why public health care expenditure might have no, or very limited, effect on the health status of the population and therefore also on health related work absence.

3.3 Results

Table 2 presents the estimation results for absence. All instrumental variable estimations were done using a two-step feasible generalized method of moments estimator, which is efficient in the presence of heteroskedacity and serial correlation.[12] First, the baseline specification (labeled IV) is presented where $eadj$, Δs_{t-1} and $\Delta(w(1-\tau^{mc}))$ were instrumented and then four other specifications are presented to serve as comparisons. In the OLS specification all regressors were treated as exogenous and in the IV-small specification the education and demographic variables were left out. In the IV-e specification e was included instead of its age-adjusted version and the IV-static specification is a static version of the first one. The omitted education- and age-groups are $El.School$ and $Pop1639$. The derivatives $ds^*/deadj|_{h_{t-1}}$ and $ds^*/de|_{h_{t-1}}$ indicate the long run effects of health care expenditure on absence. The next six statistics, which describe the relevance and validity of the instruments, are discussed in Appendix D, where I conclude that the instruments are valid and reasonable relevant.[13]

In a simultaneous test of whether $eadj$, Δs_{t-1} and $\Delta(w(1-\tau^{mc}))$ can be treated as exogenous, the null hypotheses of exogeneity could be rejected at the 10 percent level, which supports the use of instrumental variable estimators.[14] The education and age variables in levels and in first-differences, as well as the year-specific fixed effects were included since the null hypothesis of no effect could be rejected at the 10 percent level in group-wise F-tests. $\Delta Women$ and

[12] Greene (2003) describes the estimator in chapter 18.

[13] The estimations are based on 2,547 or 2,554 observations, since the use of lags and the first-difference transformation reduced the number of possible observations with 3*286, and since 27 (20 in the OLS specification) were lost due to lack of data on health care expenditure. To facilitate comparison, the OLS estimation was performed for the same years as the IV estimations.

[14] The endogeneity test is based on the difference of two Hansen-Sargan statistics and is robust against heteroscedasticity.

Table 2. Estimation results, first difference of absence

	IV	OLS	IV-small	IV-e	IV-static
$eadj$ or e	0.02	0.01	-0.03	0.03	0.03
	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)
Δs_{t-1}	0.26**	0.21***	0.37***	0.26**	
	(0.11)	(0.03)	(0.09)	(0.11)	
$\Delta(w(1-\tau^{mc}))$	-0.03	0.06***	-0.06	-0.03	-0.02
	(0.04)	(0.01)	(0.04)	(0.04)	(0.04)
$HighSchool$	-0.03	0.61		-0.22	0.44
	(0.85)	(0.82)		(0.88)	(0.93)
$University$	-3.04***	-3.56***		-3.13***	-4.13***
	(0.94)	(0.76)		(0.95)	(0.91)
$\Delta HighSchool$	-20.37**	-20.87**		-20.34**	-22.30**
	(9.31)	(8.88)		(9.28)	(8.96)
$\Delta University$	-48.52***	-60.11***		-48.63***	-51.91***
	(14.11)	(12.58)		(14.08)	(13.87)
$Women$	15.87***	17.16***		16.29***	19.92***
	(4.16)	(3.79)		(4.24)	(4.24)
$Pop4049$	-1.65	-1.56		-1.71	-0.40
	(2.70)	(2.71)		(2.69)	(2.93)
$Pop5054$	-1.46	-4.34		-1.42	-2.23
	(5.62)	(5.57)		(5.62)	(5.98)
$Pop5559$	12.73**	12.93**		12.70**	15.67***
	(5.52)	(5.46)		(5.51)	(5.91)
$Pop6064$	-4.07	-4.59		-4.45	-4.77
	(3.97)	(3.92)		(4.00)	(4.46)
$\Delta Pop4049$	34.24***	34.84***		34.22***	31.21***
	(11.48)	(11.51)		(11.44)	(11.99)
$\Delta Pop5054$	58.44***	63.80***		58.22***	61.87***
	(14.49)	(14.29)		(14.47)	(15.34)
$\Delta Pop5559$	52.92***	56.81***		52.92***	65.20***
	(15.50)	(14.35)		(15.51)	(14.98)
$\Delta Pop6064$	61.74***	70.09***		62.27***	81.85***
	(16.74)	(15.01)		(16.80)	(15.97)
$ds^*/deadj\|_{h_{t-1}}$	0.02	0.02	-0.05	0.04	
or $ds^*/de\|_{h_{t-1}}$	(0.04)	(0.04)	(0.05)	(0.06)	
Cragg-Donald	22.01		29.26	22.17	42.94
$eadj$ or e: Shea	0.72		0.81	0.63	0.74
Δs_{t-1}: Shea	0.06		0.08	0.06	
$\Delta(w(1\text{-}\tau^{mc}))$: Shea	0.09		0.09	0.09	0.09
Hansen J	0.46		0.43	0.71	0.97
Serial corr. 2	0.62	0.86	0.97	0.64	0.21
Adj. R^2	0.74	0.74	0.72	0.74	0.73
Sample size	2547	2554	2547	2547	2547

The regressions include year specific effects. Robust standard errors are shown in parentheses. The Asterisks ***, ** and * denote significance at the 1, 5 and 10 percent levels.

variables describing the share of the work force in various sectors were not included in the final regressions since these variables had no statistically significant effects. That $\Delta Women$ had no effect is surprising since women in Sweden are known to be absent more than men, but this is probably explained by low variation in gender-composition over time. Including $\Delta((w-B)(1-\tau^{mc}))$ instead of $\Delta(w(1-\tau^{mc}))$, including lagged values of $eadj$, or estimating with two-stage least squares instead of using the generalized method of moments estimators, did not change the general results.[15,16]

Table 2 shows that health care expenditure had no statistically significant effect on absence and the estimated standard errors are small enough to rule out all but a minimal effect.[17] The difference in the estimated coefficients between in the IV estimations and the OLS specifications are negligible. This might be explained by absence being measured at the municipal level whereas health care expenditure was decided at the county level, or by the county councils' weak incentive to respond to changes in absence. Using unadjusted health care expenditure (IV-e) instead of age-adjusted (IV and others) also made little difference, perhaps because there was little heterogeneity in the changes in age-composition across counties.

The effect of health care expenditure on absence is of course heterogeneous and depends on what the money is spent on. Although the purpose of this study was to estimate the effect of aggregated public health care expenditure, i.e. to estimate the average effect, such heterogeneity might cause a problem when

[15] Robust standard errors are reported since a Pagan-Hall test indicates heteroscedasticity in all specifications. For the OLS specification, a White-Koenker test was used instead.

[16] Previous literature (e.g. Henrekson and Persson, 2004) have found statistically significant effects of current unemployment and labor force participation rates on absence. Here, national variations in these variables were captured by the year-specific fixed effects, while time-invariant heterogeneity in these variables was wiped out by the first-difference transformation. Including these variables, or their lags, directly into the model did not change the general results. Whether non-working individuals are unemployed, not part of the labor force, or absent, is probably affected by variables not included in the model, making labor-force participation and unemployment endogenous. Due to this and the difficulty of finding strong instruments for two additional endogenous regressors, those variables were not included in the final specification.

[17] For the baseline specification, the 99 percent confidence interval for health care expenditure reaches down to -0.07. In percentage terms a coefficient of -0.07 translates to that a 10 percent increase in health care expenditure would only reduce absence by approximately 0.21 percent.

estimating the effect with IV methods (Heckman et al., 2006). Here, the problem would arise if the marginal expenditure identified by the instrumental variables were non-representative in terms of their effect on absence. Different instrument variables would then result in different parameters being estimated. To judge whether this is a serious problem in the present study, the baseline estimation was performed with numerous combinations of instrumental variables, which all gave similar results.[18]

The small effect of health care expenditure on absence might be explained partly by moral hazard problems; that is, individuals might reduce their personal investments in health, when public health care expenditure is increased. For example, people might exercise less and eat more unhealthy food when they have access to better health care. These moral hazard problems can also be one explanation to why several previous studies (see e.g. Aakvik and Holmås, 2006, or Nixon and Ulmann, 2006) have found no or limited effect of health care expenditure on the health statues of the population. It may also be that variations in health care expenditure in industrialized countries such as Sweden have less to do with curing and more to do with caring (Newhouse, 1977). That is, health care expenditure on the margin might be spent so that the patients' comforts increase but without leading to quicker recoveries. Other contributing explanations could be migration of sick individuals to counties with higher health care expenditure and vice versa, low efficiency in public health care, or perhaps weak correlation between total health care expenditure and that directed to the working age population, or even a weak connection between health and absence.

[18] Some instrument combinations were found to be weak or invalid (the criteria used here were Cragg-Donald>10 and Shea>0.04 for all endogenous regressors, and Hansen J>0.10), but 30 reasonably good combinations remained, including using changes in Herfindal-Indexes of political fragmentation instead of $\Delta PolmajM$ and $\Delta PolmajC$, and including additional instruments such as Δs_{t-3}. In one case, when $eadj_{t-1}$, $\Delta PolmajM$, $\Delta PolmajC$, $\Delta SocM$ and $\Delta SocC$, were replaced by a variable describing the counties' financial resources, the lag of that variable, and a variable describing the share of Health Care Party members in the county government in the prior year, the coefficient for $eadj$ was negative and statistically significant at the 10 percent level, (coeff.= -0.06, std.err=0.03). In the other cases the general results held.

Since Arellano (1989) recommended using levels instead of differences as instruments for the lagged dependent variable, s_{t-2} was also tested as an instrument instead of Δs_{t-2}. However, it was found to be weak so Δs_{t-2} was used instead. The final choice of instruments was based on the values for the baseline specification of the instrument statistics, discussed in Appendix D.

In all specifications lagged absence was significant at the 5 percent level, implying persistence. The estimated coefficient is lower in the OLS specification, which was expected since $\Delta\varepsilon_t$ and $\Delta\varepsilon_{t-1}$ will be negatively correlated, at least if ε_t and ε_{t-1} are uncorrelated.

The first difference of average after tax labor income was only significant in the OLS specification. A probable explanation to the positive estimate in that specification is that a reduction in absence lessens average labor income since the marginal non-absent individual likely has a lower labor income compared to the average in the non-absent population. If this relationship varies across municipalities, that could also account for the non-significant estimates in the other specifications. The limited effect of net income could also result from this variable having opposite substitution and income effects on the demand for absence. The coefficients might also be affected by the impact that net income has on absence through its effect on investments in health capital.

The results indicate that university graduates had a better absence development than others, and show that the coefficients for the differenced educational variables have the expected sign and relative size. Based on these estimates no conclusion can be drawn whether higher education was correlated with lower absence since those with higher education were exposed to fewer health risks at work, had better health and health development of other reasons, or had occupations that reduced their need for absence for a given health.[19] Of course, these variables might also capture characteristics that affect the absence of individuals belonging to other educational groups.

The share of women had a statistically significant positive effect, which was expected since absence increased much faster for women than for men during the study-period, partly caused by more psychological problems such as stress reactions and anxiety (Riksförsäkringsverket, 2004). Those 55 to 59 was according to the estimations, the age-group with the worst absence development. The coefficients for the differenced age variables all have the expected signs. That the coefficients for $\Delta Pop5559$ and $\Delta Pop6064$ are not even higher reflects the lower labor force participation rate in these age-groups.

Table A2 (in Appendix A) shows that the result regarding the effect of health care expenditure prevailed when absence were estimated separately for

[19] Grossman (2000) discusses possible explanations for the correlation between education and health on an individual level.

women and men. However, the estimates for the lagged dependent variable are not reliable because of weak instruments, especially for men, and this might also affect the other estimates in the dynamic IV specifications. The OLS estimations provide a possible explanation for the weak instruments for men's lagged absence, namely that the persistence in absence was relatively weak for men which results in the second lag of the dependent variable being a weak instrument for the first lag.

Absence because of sickness and disability were also estimated separately (Table E1. in Appendix E). Because of interaction between *Sickness* and *Disability* (discussed in Appendix E), both $\Delta Sickness_{t-1}$ and $\Delta Disability_{t-1}$ were instrumented with their lags and included in each estimation. The results do not allow us to reject the null hypotheses that health care expenditure has no effect on either *Sickness* or *Disability* and the estimated standard errors are small enough to rule out all but minimal effects. However, these results must be taken with caution since the instruments for $\Delta Sickness_{t-1}$ are weak, and since second-order serial correlation casts doubt on the validity of most instruments for the dynamic IV specifications and for the static specification for *Sickness*. (However, the p-values of the Hansen J statistic suggest that the instruments are valid for the *Sickness* specifications.)

4 Discussion

The effect of public health care expenditure on absence due to sickness or disability in Sweden was analyzed using an instrumental variable estimator for a dynamic panel model. Public health care expenditure was found to have no statistically significant effect on absence. This result is robust against changes in model specification and also held when separate estimations were conducted for women and men, and for absence due to sickness and disability. The standard errors were small enough to rule out all but a minimal effect of health care expenditure.

This result increases the likelihood that general health care is over-provided in Sweden, according to the model by Granlund (2007). However, health care aimed at reducing absence might still be under-provided. One possible explanation of the small effect on absence is that the correlation between expenditure on health care spent on the working population and total expenditure on health care was weak. It could also be that variation in health care expenditure had less to do with curing and more to do with caring, meaning that health care expenditure on the margin was spent so that the patients' comforts increase without leading to quicker recoveries (Newhouse, 1977). The Swedish counties have weak incentive to reduce absence, which supports either of these explanations. Due to these reasons, it should be stressed that it was the average effect of public health care expenditure on absence that was estimated, not the maximal (or potential) effect. Another set of explanation of the results is moral hazard problems, i.e. people might reduce their personal investments in health when public health care expenditure rises.

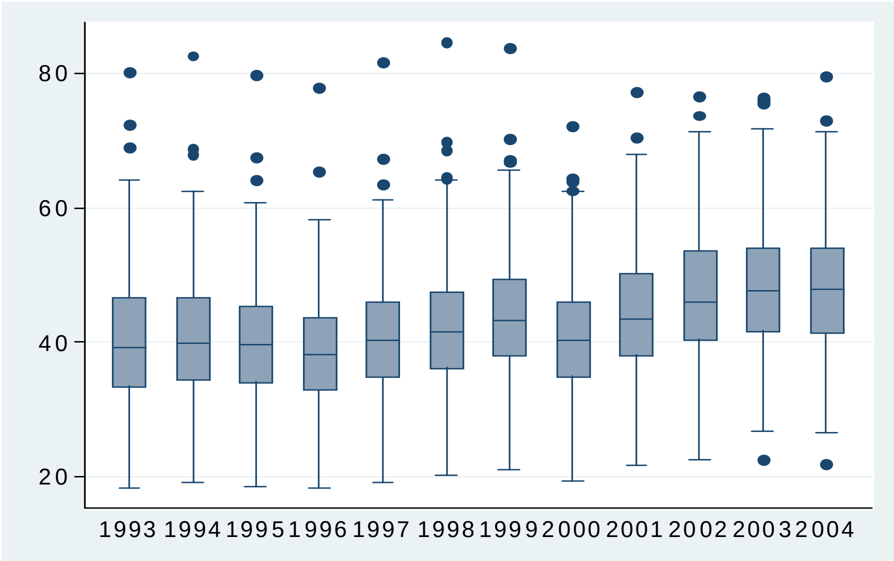
The paper relates to the literature studying the effects of access to health care or health care expenditure on health outcomes and the findings give no support for that health care expenditure on an aggregated level improves the health status of the population. However, the results from this paper might also be explained by a weak connection between health and absence. For example, the generous Swedish sickness insurance system might induce also the relatively healthy to report sick. A topic for future research could thus be to investigate the effect of health care expenditure in this sample on other health measures, e.g. on mortality rates. More research is also needed on the relationship between subcategories of health care and absence, for example, pharmaceutical expenditure. Access to health care, measured by for example waiting times, might also have an effect on absence independent of health care expenditure in general and thus be worthy of investigation in itself.

Appendix A: Tables and figures

Table A1. Data definitions and data sources.

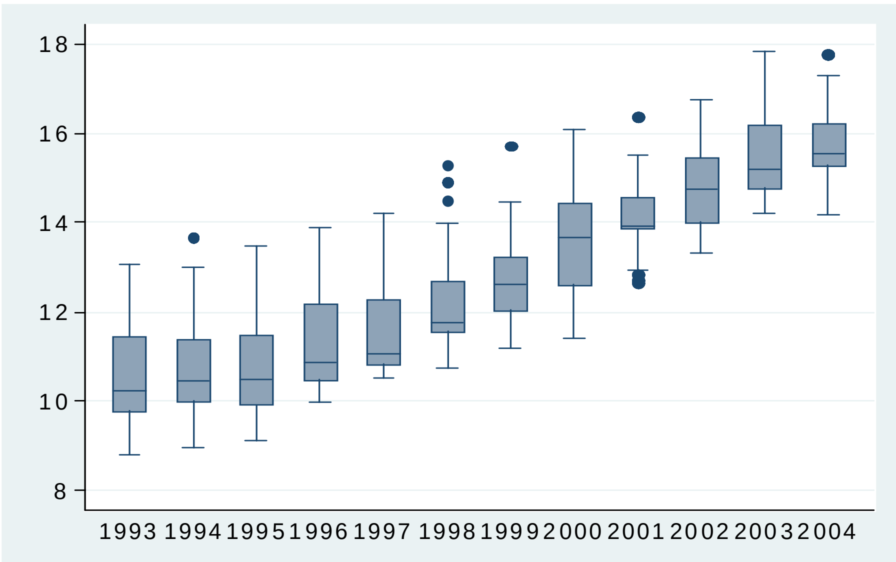
Variable	Definition	Source
Sickness	Average number of days of absence from work due to sickness for insured aged 16 to 64*	SSIA, etc.
Disability	Average number of days of absence from work due to early retirement pension/disability for insured aged 16 to 64	SSIA
Rehab	Average number of days of absence from work due to rehabilitation for insured aged 16 to 64	SSIA
s	Absence: sum of *Sickness*, *Disability* and *Rehab**	SSIA, etc.
s^{women}	Absence for women	SSIA, etc.
s^{men}	Absence for men	SSIA, etc.
e	Non-dental, non-pharmaceutical, public operating cost for health care, thousands of SEK** per capita	FCC
eadj	Age-adjusted version of e (see Appendix C)	FCC, etc.
w	[Average income from work for those aged 16 to 64 (excluding sickness and disability benefits), thousands of SEK**] / [1-absence rate], where absence rate=s/365* (insured aged 16 to 64)/(population aged 16 to 64)	SCB,SSIA
τ^M	Proportional municipality income tax rate	SCB
τ^C	Proportional county income tax rate	SCB
τ^{MC}	Sum of τ^M and τ^C	SCB
Women	Share of women in the population aged 16 to 64	SCB
*Pop*1639	Share aged 16 to 39 of the population aged 16 to 64	SCB
Pop....	Pop4049-Pop6064 have corresponding definitions	SCB
El.School	Share of the population aged 16 to 64 who's highest education was elementary school	SCB
HighSchool	Share of the population aged 16 to 64 who's highest education was high school or less than tree years after high school	SCB
University	Share of the population aged 16 to 64 with three years or longer education after high school	SCB
SocM	Share of Social Democrats and Left Party members in municipal government	SCB
SocC	Share of Social Democrats and Left Party members in county government	SCB
PolmajM	Dummy variable which takes the value 1 if either of the two traditional Swedish political blocks has a majority in the municipality government***	SCB
PolmajC	Dummy variable which takes the value 1 if either of the two traditional Swedish political blocks has a majority in the county government***	SCB

The data sources are Swedish Social Insurance Agency, SSIA, The Federation of Swedish County Councils, FCC, and Statistics Sweden, SCB.
All monetary variables are deflated by CPI and expressed in 2004 years prices.
*Only absence from the 15th day of each spell is included.
**On 21 December 2006, USD/SEK=6.81.
***The two political blocks consist of the Social Democratic Party and the Left Party; and the Moderate Party, the Liberal Party, the Christian Democrats and the Centre Party.



Note: The boxes include observations from the 25th to 75th percentiles. The whiskers are 1.5 times the length of the boxes or equal the distance from the box to the minimum or maximum values, whichever is smallest. The dots indicate outside values.

Figure A1. Box plot for absence, *s*



See note to Figure A1.

Figure A2. Box plot for age-adjusted health care expenditure, *eadj*

Table A2. Estimation results, first-difference of women's and men's absences

	Women			Men		
	IV	OLS	IV-static	IV	OLS	IV-static
$eadj$	0.05	0.05	0.06	-0.04	-0.01	-0.04
	(0.05)	(0.04)	(0.06)	(0.04)	(0.03)	(0.04)
$\Delta s^{\dagger}_{t-1}$	0.08	0.19***		0.20	0.08**	
	(0.12)	(0.03)		(0.22)	(0.03)	
$\Delta(w(1-\tau^{mc}))$	-0.04	0.07***	-0.04	-0.02	0.03**	-0.01
	(0.05)	(0.02)	(0.05)	(0.04)	(0.01)	(0.03)
$HighSchool^{\dagger}$	3.26*	2.91**	3.76**	-1.17	-1.18*	-1.27*
	(1.68)	(1.45)	(1.56)	(0.72)	(0.70)	(0.76)
$University^{\dagger}$	-4.47***	-4.66***	-4.77***	-0.47	-1.04*	-0.76
	(1.13)	(0.99)	(1.12)	(0.66)	(0.54)	(0.59)
$\Delta HighSchool^{\dagger}$	-16.81	-18.62*	-17.20*	-7.91	-8.11	-11.10
	(10.40)	(10.31)	(10.45)	(8.19)	(7.29)	(7.36)
$\Delta University^{\dagger}$	-12.74	-28.09*	-12.32	-47.24***	-55.13***	-54.77***
	(16.89)	(15.17)	(16.87)	(14.44)	(11.01)	(11.85)
$Pop4049^{\dagger}$	-0.23	-1.19	0.13	-4.34*	-4.02	-4.15
	(3.68)	(3.51)	(3.78)	(2.49)	(2.55)	(2.61)
$Pop5054^{\dagger}$	-3.94	-5.68	-3.91	5.23	4.34	5.63
	(5.85)	(5.59)	(6.05)	(4.53)	(4.53)	(4.61)
$Pop5559^{\dagger}$	28.26***	24.56***	30.16***	0.83	0.39	0.20
	(6.39)	(5.79)	(6.29)	(4.69)	(4.64)	(4.68)
$Pop6064^{\dagger}$	-11.14**	-11.60**	-11.70**	0.47	-0.34	1.53
	(5.40)	(5.00)	(5.66)	(3.94)	(3.97)	(4.06)
$\Delta Pop4049^{\dagger}$	24.63*	27.43**	24.23*	30.76***	28.39***	28.36***
	(13.53)	(13.50)	(13.76)	(9.91)	(9.71)	(9.75)
$\Delta Pop5054^{\dagger}$	50.42***	52.40***	52.67***	29.39**	29.53**	26.76**
	(16.58)	(16.38)	(16.90)	(12.60)	(12.08)	(12.26)
$\Delta Pop5559^{\dagger}$	44.81**	46.97***	48.40***	32.29**	38.13***	40.60***
	(17.92)	(17.38)	(17.93)	(14.90)	(12.01)	(12.16)
$\Delta Pop6064^{\dagger}$	80.05***	80.77***	85.63***	62.84***	72.70***	75.11***
	(18.03)	(17.31)	(17.84)	(18.31)	(12.72)	(13.09)
$ds^{*}/deadj\|^{\dagger}_{h_{t-1}}$	0.06	0.07		-0.05	-0.01	
	(0.06)	(0.05)		(0.05)	(0.04)	
Cragg-Donald	14.54		42.33	6.02		44.81
$eadj$ or e: Shea	0.72		0.74	0.75		0.75
$\Delta s^{\dagger}_{t-1}$: Shea	0.04			0.02		
$\Delta(w(1\text{-}\tau^{mc}))$:Shea	0.10		0.10	0.11		0.11
Hansen J	0.88		0.93	0.90		0.80
Serial corr. 2	0.18	0.94	0.25	0.31	0.26	0.34
Adj. R^2	0.67	0.68	0.67	0.62	0.63	0.63
Sample size	2261	2268	2261	2261	2268	2261

† indicates that the variable is gender-specific. Also, see notes to Table 2.

Appendix B: Missing data and changes in the absence variables

Data on the absence variables were missing for some of the municipalities in 1997, 1998 and 1999. Instead, aggregated data were reported for two (or sometimes three or four) municipalities in the same county. For six municipalities, data for all three years were missing, and instead the aggregated data for three pairs of municipalities were reported. For 37 municipalities, data for 1998 and 1999 were missing, and instead the aggregated data for 12 pairs of municipalities, and three groups of three, and one group of four municipalities were reported. For seven municipalities, data for 1999 were missing, and instead the aggregated data for one group of three and one group of four municipalities were reported. Thus in total, 99 observations lacked data on the absence variables.

The absolute number of days of absence (*days*) as well the number of insured for each group of municipalities is known. I assumed that the *days* were divided among the municipalities in each group in proportion to their share of *days* before and after the years of missing data. That is

$$days_{iM} = \frac{days_{iB} + days_{iT}}{days_{pB} + days_{pT}} days_{pM},$$

where i indicates the ith municipality and p indicates the group of municipalities to which it belongs. B indicates the last year before, T the first year after, and M the years of missing data. This assumption was sufficient to get an estimate of *days* for each municipality for which data were missing for only one year. The estimated *days* was then divided by the number of insured, estimated in the same way, to obtain an estimate of s for each municipality.

Using the maximum entropy method (Wilson, 1970), *days* for those lacking data for two or three years was estimated as

$$days_{it} = \frac{days_{iM} * days_{pt}}{days_{pM}}.$$

The number of insured was estimated in the same way, and then an estimate of s was calculated for each municipality each year. *Sickness*, *Disability*, and *Rehab* were each estimated in this way, which is the least biased estimate possible with the information given (Jaynes, 1957).

For 1999, data on the absence variables were missing, and instead data for October 1998 to September 1999 were reported. I used the numbers reported inflated with 1/8 of the change from 1998 to 2000, which assumes that changes for the fourth quarter were of the same magnitude as the average change for the other three quarters from 1998 to 2000, and that half of this change occurred each year.

Estimations excluding the observations with missing data, and using the average of the values for 1998 and 2000 instead of the calculated values for 1999, give the same general results for the baseline specification.

As mentioned earlier, the absence variables include only days compensated by the Swedish Social Insurance Agency, not days compensated by employers. During 1992-1996 and April 1998-June 2003 employees were compensated by the employer for the first 14 days of absence; during January 1997-March 1998 for the first 28 days; and during July 2003-December 2004 for the first 21 days.[20] Following Henrekson and Persson (2004), this was addressed using data from The Confederation of Swedish Enterprise.

Their data cover a reasonably representative sample of 2,500 private sector establishments and 220,000 employees. Absence from work due to sickness was categorized by the length of the absence spells and separate figures were reported for nine Swedish regions.[21] Assuming that the absence pattern was the same for all municipalities belonging to the same region, and the same for the public sector as for the private sector, the original data for $Sickness$ (and s) were adjusted to give absences from the 15th day of absence for all years. This was done by multiplying the original $Sickness$-variable by the percentage of work time lost due to sickness absence from the 15th day of absence divided by the percentage of work time lost due to sickness that was covered by the Swedish Social Insurance Agency.[22,23] The variable s was then created by adding

[20] The self-employed were allowed to choose among insurance plans which differed in when they began to reimburse for lost income due to sickness. One plan stipulated that the self-employed were reimbursed first from the 31st day, but this should have a small effect, if any, on the estimations.

[21] The data for 1997 and 1998 were published in SAF (1998) and SAF (1999), respectively. Data for 2003 and 2004 were provided directly by the company, Löneanalyser AB.

[22] From the data it is not possible to directly identify the shares of work-hours lost due to absence during days 15 to 21; instead the shares for days 15 to 20 are reported, which were multiplied by 7/6. Unfortunately this causes a slight overestimation of absence for the years 2003 and 2004, because of less absence the 21st day compared to the average for days 15-20. The absence patterns for the first and second halves of 2003 were assumed identical in order to adjust the data.

[23] For 1997 $Sickness$ was multiplied by a factor ranging from 1.12 to 1.23 depending on the

Disability and *Rehab* to the adjusted version of *Sickness*. s^{women} and s^{men} were created similarly. Table 1 reports adjusted versions of these variables.

Appendix C: Age-adjustment of health care expenditure

The age-adjustment was based on an index of age dependent health care consumption produced by Statistics Sweden, which was used to calculate intergovernmental equalization grants for counties until 1995.[24]

For somatic short-term health care, psychiatric care, and geriatric care, the national average of treatment days for the age-groups 0-14, 15-44, 45-64, 65-79 and 80- were provided each year by the National Board on Health and Welfare.

The average number of physician consultations in primary health care and hospital connected health care were provided by the County Council of Skåne for 1991-1993 (covering only the county of Malmöhus, which became a part of Skåne in 1998) and for 2001-2004. To avoid regional differences, only data from former Malmöhus were used for the last period as well. These sources were used since no national figures were available. For 1991-1993 the data were reported for the same five age-groups as above, but for 2001-2004 eight age-groups were used: 0-4, 5-14, 15-24, 25-44, 45-64, 65-74, 75-84, and 85-. Based on this data, percentage changes each year were estimated for the two consultations types for the age-groups 0-14, 15-44, 45-64, and 65-. Then the number of consultations in each of the eight age-groups were calculated for 1993-2000 using the estimated percentage changes for the group 0-14 for the age-groups 0-4 and 5-14, etc.

Using these figures and population data, the expected number of treatment days and physician consultations per inhabitant were calculated for each county, each year, for each type of health care. These numbers were then divided by the national average to obtain indexes, which were aggregated using each health care category's national cost-share each year as weight. The cost-shares were

region. The corresponding figures for 1998, 2003, and 2004 are 1.03 to 1.05; 1.01 to 1.04; and 1.01 to 1.07, respectively. For all other years the factor was of course 1.

[24] Two other methods have been used by Statistics Sweden since 1995, the first based on a regression of health care expenditure on a number of macro-variables. This method was not used here since it can capture not only differences in need for health care but also differences in preference and resources. The second method was not used here since it depends on the number of patients with particular diagnoses, which is likely to be endogenous.

calculated from data obtained from Statistics Sweden for the years 1993-2003. The average cost-shares were; somatic short-term health care (0.5); psychiatric care (0.08); geriatric care (0.05); physician consultations in primary health care (0.19) and physician consultations in hospital connected health care (0.18). These five categories accounted for approximately 95 percent of non-dental non-pharmaceutical health care consumption. Finally, age-adjusted health care expenditure (*eadj*) was calculated by dividing health care expenditure (e) with the appropriate index for each observation.

Appendix D: Relevance and validity of the instruments

Table 2 reports the Cragg-Donald weak identification statistic, which is the smallest eigenvalue of the matrix analog to the F-statistic from the first-stage regressions. Since the models include several endogenous variables, this statistic is reported instead of the F-statistic. Based on the tabulation by Stock and Yogo (2005) of critical values for the Cragg-Donald weak identification statistic, the instrument is judged to be strong if the statistic is above 13.95 in the dynamic or 15.72 in the static specification.[25] The statistic is well above these values in all specifications.

As a complement to this test, Shea's (1997) partial R-squared measure of instrument relevance for models with multiple endogenous variables is reported. Shea did not provide any critical values, but mentioned 0.05 as an example when the instrument set is not very relevant for an endogenous regressor, and noted that low relevance increases asymptotic standard errors and increases the inconsistency of the estimates whenever instruments are not perfectly exogenous. The Shea values in Table 2 are above 0.05 for all endogenous regressors, especially for *eadj*, but are quite close (0.06) for Δs_{t-1} in two specifications. Together the Cragg-Donald and Shea statistics indicate that the instruments

[25] The critical values used are those for a maximum bias of 0.05 relative to OLS for a two-stage least squares (TSLS) estimator. No critical values are provided for a general means of moments (GMM) estimator, and these values are only approximate for the GMM estimator, since it diverges somewhat from the TSLS estimator because of heteroscedasticity and autocorrelation. But at least the test indicates strong instruments for the TSLS estimator, which, as discussed in the text, gives similar results.

are, at least, reasonably relevant for all the endogenous regressors.

The Hansen J statistic is the p-value of the Hansen test of overidentifying restrictions, where the joint null hypothesis is that the instruments are valid, i.e., uncorrelated with the error term. This test is consistent in the presence of heteroskedasticity and serial correlation, and supports the exogeneity of the instruments used.

Serial corr. 2 reports the p-value of a t-test of serial correlation of the second order. This test was conducted since the exogeneity assumptions for two of the four instruments were based on the assumption of no second-order serial correlation. The test can be viewed as complementary to the Hansen J test, which would also indicate that the instruments were invalid if second-order serial correlation were too strong. For all specifications, no statistically significant second-order serial correlation was found. Thus, both Hansen J and the Serial corr. 2 test support the assumption that the instruments are valid for all specifications reported in Table 2.

Appendix E: Sickness and disability

The basic transitions between the three mutually exclusive states that an individual can occupy are illustrated in Figure E1.[26] The illustrated flows are those that are driven by other factors than economic incentives, health status and work characteristics. The vertical arrows illustrate the effects of habit formation. The diagonal arrow from $Sickness_{t-1}$ to $Disability_t$ illustrates the flow caused by information increasing over time regarding the expected duration of individual's reduced work capacity. That is, this arrow shows the flow between the two states caused by better predictions about the individual's future work capacity, conditioned on the observed present one.

The flow from $Work_{t-1}$ to $Sickness_t$ illustrates the inflow to absence that is affected by previous absence in the municipality, through its effect on the disutility of absence.[27] One possibility is that this habit formation is not specific

[26] *Rehab* is of minor importance (Table 1) so no separate analysis was done with it as dependent variable.

[27] It is less likely that there is a direct flow from $Work$ to *Disability* of this reason and no such arrow is therefore drawn. It is however possible, that for example individuals with long-lasting small reduction in work capacity go directly from $Work$ to *Disability*, as a consequence of external habit formation.

to the type of non-work state. For example, a high rate of people on disability pension in a municipality might also reduce the social cost of being on sick leave, resulting in a correlation between $Disability_{t-1}$ and $Sickness_t$, even if no one actually moved in this direction, simply because a high rate of *Disability* caused a flow from $Work$ to $Sickness$. Table E1 presents the estimation results for *Sickness* and *Disability*, which are discussed briefly in the text.

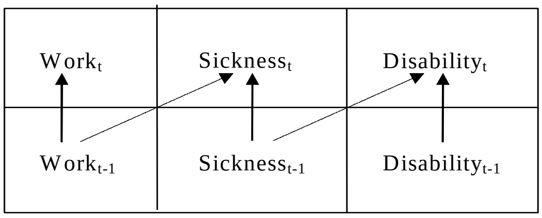


Figure E1. Interactions between $Work$, *Sickness*, and *Disability*

Table E1. Estimation results, first difference of *Sickness* & *Disability*

	Sickness			*Disability*		
	IV	OLS	IV-static	IV	OLS	IV-static
$eadj$	0.05 (0.04)	0.04* (0.02)	0.03 (0.03)	-0.04 (0.03)	-0.01 (0.02)	0.01 (0.03)
$\Delta Sick._{t-1}$	-0.23 (0.16)	0.10*** (0.03)		0.53*** (0.13)	0.05** (0.02)	
$\Delta Dis._{t-1}$	-0.22** (0.11)	-0.10*** (0.03)		0.47*** (0.09)	0.34*** (0.03)	
$\Delta(w(1-\tau^{mc}))$	-0.04 (0.04)	0.00 (0.01)	-0.05 (0.03)	0.01 (0.03)	0.06*** (0.01)	0.03 (0.03)
$HighSchool$	1.62* (0.90)	1.07 (0.72)	1.14 (0.77)	-1.53** (0.73)	-0.33 (0.63)	-0.57 (0.74)
$University$	-0.87 (1.00)	-0.46 (0.70)	0.10 (0.76)	-1.90** (0.83)	-3.08*** (0.61)	-4.15*** (0.72)
$\Delta HighSchool$	3.83 (8.14)	5.09 (8.01)	5.90 (7.99)	-17.05** (7.50)	-20.64*** (6.99)	-22.68*** (6.87)
$\Delta University$	-20.13 (12.75)	-22.48* (11.51)	-15.84 (11.98)	-19.93* (11.77)	-31.62*** (9.62)	-31.14*** (10.49)
$Women$	10.17** (4.43)	8.51** (3.44)	6.72* (3.60)	5.24 (3.47)	7.62*** (2.86)	12.47*** (3.31)
$Pop4049$	1.09 (3.00)	-0.66 (2.58)	-0.32 (2.72)	-2.71 (2.27)	-0.96 (1.99)	-0.40 (2.24)
$Pop5054$	-4.99 (6.10)	-3.15 (5.07)	-3.12 (5.45)	2.24 (5.21)	-1.25 (4.19)	-0.02 (4.63)
$Pop5559$	14.04** (5.89)	10.22** (5.19)	11.54** (5.37)	0.86 (4.93)	2.79 (4.38)	4.78 (4.88)
$Pop6064$	0.01 (4.61)	-0.87 (3.96)	-0.01 (4.06)	-5.91* (3.54)	-3.54 (3.13)	-5.40 (3.48)
$\Delta Pop4049$	13.38 (11.98)	17.93* (10.49)	17.22 (10.86)	21.35** (10.57)	15.54* (9.04)	13.44 (9.41)
$\Delta Pop5054$	36.49** (14.61)	33.00** (13.05)	32.47** (13.42)	20.56* (11.96)	25.87** (10.70)	27.57** (11.50)
$\Delta Pop5559$	35.74** (15.92)	24.99* (13.99)	24.56* (13.97)	19.53 (13.72)	32.26*** (11.63)	40.99*** (11.79)
$\Delta Pop6064$	20.10 (17.39)	9.19 (14.31)	2.28 (14.31)	42.71*** (14.05)	58.25*** (11.12)	77.57*** (11.87)
$\mathrm{d}s^*/\mathrm{d}eadj\|_{h_{t-1}}^{\boxtimes}$	0.04 (0.03)	0.05* (0.03)		-0.07 (0.06)	-0.02 (0.03)	
Cragg-Donald	9.92		42.94	9.92		42.94
$eadj$ or e: Shea	0.69		0.74	0.69		0.74
$\Delta Sick._{t-1}$:Shea	0.03			0.03		
$\Delta Dis._{t-1}$:Shea	0.09			0.09		
$\Delta(w(1\text{-}\tau^{mc}))$:Shea	0.09		0.09	0.09		0.09
Hansen J	0.27		0.37	0.00		0.44
Serial corr. 2	0.00	0.02	0.02	0.00	0.00	0.52
Adj. R^2	0.59	0.62	0.61	0.70	0.77	0.74
Sample size	2547	2554	2547	2547	2554	2547

$^{\boxtimes}$ $\mathrm{d}s^*/\mathrm{d}eadj|_{h_{t-1}}$ states the long run effect of *eadj* on *Sickness* and *Disability*, respectively excluding the effect that goes through the other variable. Also, see notes in Table 2.

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Fixed budgets as a cost containment measure for pharmaceuticals

David Granlund, Niklas Rudholm and Magnus Wikström

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Consumer Information and Pharmaceutical Prices: Theory and Evidence*

David Granlund† and Niklas Rudholm‡

†Department of Economics, Umeå University,
SE-901 87 Umeå, Sweden

‡The Swedish Retail Institute (HUI),
SE-103 29 Stockholm, Sweden; and
Department of Economics, University of Gävle,
SE-801 76 Gävle, Sweden.

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Abstract

In this paper, the impact of increased consumer information on brand name and generic pharmaceutical prices is analyzed both theoretically and empirically. The theoretical results show that an increase in information is likely to reduce the price of brand name pharmaceuticals, while the results regarding generics are less clear. In the empirical part of the paper, the introduction of the substitution reform in the Swedish pharmaceuticals market in October 2002 is used as a natural experiment regarding the effects of increased consumer information on pharmaceutical prices. The results clearly show that the reform has lowered the price of both brand name- and generic pharmaceuticals.

Key Words: pharmaceutical industry; generic competition; generic drugs; brand name drugs

JEL classification: D80; D83; L65; I11

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1 Introduction

In this paper, a reform of the Swedish pharmaceutical market will be used to study the effects of increased consumer information about price differences in medically equivalent treatments (branded and generic drugs) on the pricing policy of pharmaceutical firms.

There is a vast theoretical literature concerning the effects of imperfect consumer information on pricing and market structure. Stigler (1961) showed that imperfect information creates market power and prices above competitive levels; while Diamond (1971) showed that if information is costly, this could lead to an equilibrium where firms charged the monopoly price rather than the competitive price. In addition, Salop and Stiglitz (1977) presented a model where low-cost stores had higher sales because low search cost individuals actively seek them out, while only high search cost individuals patronized the high cost stores. They also established that costly information could lead to market equilibria where the law of one price does not hold. In a related study, Stiglitz (1979) showed that under imperfect information, increasing the number of firms could actually reduce competition and increase prices.

Turning to theoretical models of pharmaceutical markets, Frank and Salkever (1992) have shown that an increase in the number of informed consumers will lead to lower brand name pharmaceutical prices. Also, based on a conjecture by Grabowski and Vernon (1992) that an increase in the number of informed customers (i.e. cross-price sensitive consumers) would increase the negative impact of entry on brand name pharmaceutical prices, Frank and Salkever showed that, theoretically, this is not necessarily the case. Our theoretical model, presented in section 3 below, will be a modified version of the Frank and Salkever model.

Only a few empirical tests of how consumer information affects prices in health care markets exist. Pauly and Satterthwaite (1981) studied how the market for a reputation good (i.e. a good that is marketed mainly through recommendation from friends, relatives or colleagues) such as physician services was affected by additional primary care physicians entering the market. Since this increases consumer search costs in their model, this will also make each primary care physicians demand curve become less elastic, leading to higher prices. In a more recent paper, Sorensen (2000) studied the relationship between imperfect consumer information and prices among prescription pharmaceuticals. The data were collected from pharmacies in upstate New York, and the price differences among medically equivalent prescriptions were found to be large. In fact, on average the highest listed price exceeded the cheapest available alternative by as much as 50 percent. In addition, the results give support to a consumer search cost model, since frequently purchased pharmaceuticals had both lower markups and lower price differences when compared to one-time prescription pharmaceuticals.

In October 2002, the Swedish pharmaceuticals market was reformed. The reform (Lag 2002:160) required that pharmacists inform the consumers if there are substitute products available, as well as that the cheapest available substitute product would be provided within the Swedish pharmaceuticals insurance system. The reform also required that the consumers are given the opportunity to buy the prescribed pharmaceutical product instead of the cheapest available product, paying the difference in price between the products themselves. This means that under the new regulations, consumers have more information about the price difference between the prescribed (often brand name) product and the cheapest available (generic) alternative than they had prior to the reform. As such, the introduction of the substitution reform can be seen as a natural experiment where consumers are given more information about price differences between identical medical treatments.

In this paper, the introduction of the substitution reform will be used to empirically measure the effects of increased consumer information on brand name and generic pharmaceutical prices. Our theoretical model shows that an increase in information is likely to reduce the price of brand name pharmaceuticals, but the results regarding generics are less clear. Hence, the main hypothesis to be tested is if the substitution reform, by increasing consumer information about pharmaceutical prices and available generic substitutes, has decreased the price of brand name and/or generic pharmaceuticals. In addition, we will test whether the possible price response differs between brand name and generic drugs and also study additional heterogeneity in the reform-effect, suggested by the theoretical model. The empirical analysis is performed using monthly data on pharmaceuticals sold January 2001 to October 2006.

The paper contributes to the existing literature in the following ways; first, contrary to previous theoretical studies, our theoretical model analyses how both brand name- and generic pharmaceuticals are affected by increased consumer information. Second, previous studies of the introduction of substitution reforms in European pharmaceuticals markets (e.g. Buzzelli et al., 2006) use pharmaceutical price indexes as their dependent variable. As such, they cannot discriminate between those changes in the pharmaceutical price index caused by changed pharmaceutical prices and those caused by changed quantities for brand name- and generic drugs. The use of pharmaceutical price indexes also makes it impossible to study heterogeneity in the reform-effect among brand name- and generic drugs. The introduction of the substitution reform and the use of individual pharmaceutical price data in our paper make it possible to study both how pharmaceutical prices were affected by the reform, as well as possible heterogeneity in the reform-effect among brand name pharmaceuticals and generics. The main finding from this paper is that the reform has lowered the price of both brand name- and generic pharmaceuticals.

The next section describes the substitution reform, while section 3 presents the theoretical model, based on the Frank and Salkever model. Section 4 presents the data and the empirical model to be used in this study, as well as the results from the estimation of the empirical model. Finally, in section 5 the paper's conclusions are presented.

2 The substitution reform

The substitution reform came into effect on October 1, 2002. As mentioned above, the reform required that pharmacists inform the consumers if there are substitute products available, as well as that the cheapest available generic substitute product (which is considered to be a perfect substitute for the brand name drug by the Swedish Medical Products Agency, SMPA) would be provided within the Swedish pharmaceuticals insurance system. The pharmacist must also inform the consumers that they can buy the prescribed pharmaceutical product instead of the generic product if they pay the difference in price between the products themselves. As such, the new regulations provide consumers with more information about the price difference between the prescribed product and the cheapest available generic alternative then before. Finally, the reform requires that pharmacists substitute the prescribed pharmaceutical product to the cheapest available generic (or parallel imported product) in cases when neither the prescribing physician prohibits the switch for medical reasons, nor the consumer chooses to pay the price difference between the prescribed and the generic alternative. In cases where the physician prohibits the switch due to medical reasons the consumer is still reimbursed.[1]

The reform was supposed to lower pharmaceutical costs in two different ways; directly, as more expensive pharmaceuticals were exchanged for cheaper generic copies, and indirectly through increased price competition. The latter effect will be studied in this paper using data from the county of Västerbotten, Sweden. Pharmaceuticals are sold through a nation wide government owned monopoly, the National Corporation of Swedish Pharmacies (NCSP), which is required to charge a uniform price for each pharmaceutical product in Sweden. In addition, for a product to be included in the Swedish pharmaceuticals insurance system the price charged by the pharmaceutical firms from the NCSP has to be authorized by the Pharmaceutical Benefits Board (PBB).

Before the substitution reform, a reference price system introduced in January 1993 was in effect. Under that system, the Swedish National Social Insurance Board (SNIB) set a reference price equal to 110 percent of the price of the cheapest available generic product, and all costs exceeding this reference price were to be borne by the consumer (RFFS 1992:20, 1996:31). The effects of the reference price system on pharmaceutical prices have been analyzed previously (see e.g. Aronsson et al., 2001; Bergman and Rudholm, 2003). The results from Aronsson et al. showed that the introduction of the reference price system reduced brand name pharmaceutical prices for pharmaceuticals where generic substitutes were available at the introduction of the system. However, Bergman and Rudholm found that the Swedish reference price system had been effective only for those products which already faced generic competition at the introduction of the system. Before the system was introduced, generic entrants typically set prices far below the brand name product, without capturing more than a relatively small fraction of the market (Aronsson et al.). When the system was introduced, the brand name manufacturers were forced to lower their prices to a level close to that of the generics, in order for consumers to be fully reimbursed. However, after the system had been introduced, generic entrants had incentives to set prices in order to stay within the 10 percent limit set by the reference price system (Bergman and Rudholm).

[1]During the first 15 months after the reform, physicians chose to deny the exchange in 3 percent of the cases (National Corporation of Swedish Pharmacies et al., 2004).

3 Theoretical model

To study the effects of increased information on brand name and generic drugs, we turn to the model developed by Frank and Salkever (1992). We modify their model to be able to analyze the price of generics as well. In a specific market there is one brand name firm, b, and $n-1$ generic firms, that all produce one pharmaceutical product each and have identical cost functions. Since there are consumers who are willing to pay the difference in price between the brand name and generic pharmaceutical in order to get the brand name pharmaceutical, it will be assumed that consumers actually view the brand name and the different generic drugs as close substitutes, instead of perfect ones.[2] Therefore, a firm cannot get all informed costumers by setting their price slightly below the others. The consumers only differ in their preferences for the drugs and the information they have about prices.

Several studies have reported that brand name products, in spite of considerably higher prices than the available generic substitutes, are able to maintain dominant positions in terms of market shares (e.g. Bond and Lean, 1977; Statman, 1981; Hurwitz and Caves, 1988). For the Swedish pharmaceuticals market, Aronsson et al. (2001) reported that some of the products in their study maintained market shares in excess of 80 percent even though they had a more than 50 percent markup over the price of the available generic substitutes. As such, the principal difference between a generic firm and the brand name firm in our theoretical model is that the brand name firm will be assumed to have a higher market share compared to the generic firm at equal prices, due to consumer preferences for the brand name product. Due to its high market share, we assume that a change in the brand name price will have a significant impact on the generic prices and treat the brand name firm as a Stackelberg leader which incorporates the generic firms' price-responses in its pricing decision.[3] It will also be assumed that the generic firms take all other prices as given in the sense that they do not take into account how their pricing decisions will affect that of other pharmaceutical firms. As such, the analysis will start by analyzing the pricing decision made by the generic firms.

[2] During the first 15 months after the reform, consumers chose to deny the exchange in approximately 5 percent of the cases (Apoteket AB et al., 2004).

[3] The qualitative results hold even if the brand name firm is treated as a Nash player towards the generic ones.

The demand function for generic product i is written

$$Q_i = \alpha D_i^I(P_i, P_j, P_b, n) + (1-\alpha) D_i^{UI}(P_i, n). \tag{1}$$

α is the share of informed consumers in the market, who know the prices of all available pharmaceutical products, and $(1-\alpha)$ is the share of uninformed consumers in the market, who only know the price of the drug they are buying. αD_i^I and $(1-\alpha)D_i^{UI}$ represent the demand facing the generic firm i from informed and uninformed consumers, respectively. P_i, P_j and P_b are the prices of the generic drug i, all other generic drugs and the brand name drug, respectively, which together with the number of pharmaceutical products, n, affect the demand for the generic drug i from the informed consumers. The demand from the uninformed consumers is not affected by other prices, but by the number of pharmaceutical products.

The generic firm i's profit function is written

$$\begin{aligned} \pi_i &= P_i * [\alpha D_i^I(P_i, P_j, P_b, n) + (1-\alpha) D_i^{UI}(P_i, n)] \\ &\quad - C_i(\alpha D_i^I(P_i, P_j, P_b, n) + (1-\alpha) D_i^{UI}(P_i, n)), \end{aligned} \tag{2}$$

where $C_i(Q_i)$ is the firm's cost function. The marginal cost is assumed to be positive and constant.[4] The firm chooses P_i to maximize the profit, which gives the following first order condition

$$\begin{aligned} d\pi_i/dP_i &= \left(P_i - \frac{dC_i}{dQ_i}\right) * \left[\alpha \frac{dD_i^I}{dP_i} + (1-\alpha)\frac{dD_i^{UI}}{dP_i}\right] \\ &\quad + \alpha D_i^I(P_i, P_j, P_b, n) + (1-\alpha) D_i^{UI}(P_i, n) = 0. \end{aligned} \tag{3}$$

The price function of generic product i can thus be written $P_i(\alpha, n, P_j, P_b)$. All generic firms are assumed to face identical demand and cost functions and will therefore set the same price, which we denote by P_g. The best response function of generics is written $P_g(\alpha, n, P_b)$.

Accordingly, the brand name producer's demand function is written

$$Q_b = \alpha D_b^I(P_b, P_g(\alpha, n, P_b)) + (1-\alpha) D_i^{UI}(P_b, n), \tag{4}$$

[4] This assumption is not crucial for the qualitative results from the model, but it simplifies expressions and makes them easier to interpret.

which gives the following first order condition

$$\begin{aligned} d\pi_b/dP_b &= \left(P_b - \frac{dC_b}{dQ_b}\right) * \left[\alpha\left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g}\frac{\partial P_g}{\partial P_b}\right) + (1-\alpha)\frac{dD_b^{UI}}{dP_b}\right] \\ &\quad +\alpha D_b^I(P_b, P_g(\alpha, P_b, n)) + (1-\alpha)D_b^{UI}(P_b, n) = 0. \qquad (5) \end{aligned}$$

Equation (3) and equation (5) differ from each other since the brand name firm get higher sales compared to a generic firm with the same price and because the brand name firm act as a Stackelberg leader. If $\frac{\partial P_g}{\partial P_b} > 0$,[5] which we assume, either one of these differences guarantee that the brand name firm will set a higher price than the generic ones.

3.1 The effect of an increase in information

The effect of an increase in the share of informed consumers (i.e. an increase in α) on the price of generic drug i is written

$$\begin{aligned} dP_i/d\alpha &= \{\left(P_i - \frac{dC_i}{dQ_i}\right) * \left(\frac{dD_i^I}{dP_i} - \frac{dD_i^{UI}}{dP_i}\right) + \left(D_i^I - D_i^{UI}\right) \\ &\quad + \frac{dP_i}{dP_b}\frac{dP_b}{d\alpha} + \frac{dP_i}{dP_j}\frac{dP_j}{d\alpha}\}/(-\delta_i) \qquad (6) \end{aligned}$$

We assume that the second order sufficient condition for a maximum, $\delta_i < 0$, is fulfilled (Appendix 2). The mark-up, $P_i - \frac{dC_i}{dQ_i}$, is positive. The second term in the numerator, and therefore the first product, will be negative if the own price response of informed consumers is greater than the own price response of uninformed consumers, which is a reasonable assumption. The next two terms show how the total demand for generic i is affected by a change in the share of informed consumers. These terms are jointly positive since informed consumers are more likely to patronize the generic product i due to price information, while consumers in the uniformed group choose the pharmaceutical product closest to their personal preferences, brand name or generic.

The final terms show the indirect effect on P_i of an increase in α, working through the effect of α on other prices. The terms $\frac{dP_i}{dP_j}$ and $\frac{dP_i}{dP_b}$, which are discussed in more detail in Appendix 1, are assumed to be positive. The last two terms therefore imply that a generic firm is more likely to reduce its price

[5] See Appendix 1 for a discussion of these assumptions.

if the brand name price is reduced and that any price change of a generic firm, due to symmetry, is enhanced by the price changes of the other generic firms.

To sum up, the direct effect of an increase in the share of informed consumers on the generic prices is not clear cut. The first product works in the direction of lower prices as long as the consumer becomes more price sensitive when information about prices increases. On the other hand, the next two terms in the numerator, D_i^I and D_i^{UI}, together have a positive effect on the price change for generic product i. Finally, the sign of the indirect effect depends on the effect of α on the brand name price, to which we now will turn.

The effect of an increase in α on the brand name price is written

$$\begin{aligned} dP_b/d\alpha &= \{\left(P_b - \frac{dC_b}{dQ_b}\right) * \left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g}\frac{\partial P_g}{\partial P_b} - \frac{dD_b^{UI}}{dP_b}\right) \\ &\quad + \left(D_b^I - D_b^{UI}\right) + \frac{dP_b}{dP_g}\frac{dP_g}{d\alpha}\}/(-\delta_b). \end{aligned} \tag{7}$$

$P_b - \frac{dC_b}{dQ_b}$ is positive and larger than the mark-up for a generic firm. The second term will be negative if the own price response of informed consumers is greater than the own price response of uninformed consumers, which seems reasonable to assume. However, this term will be less negative than the corresponding one for the generic firms due to the brand name firm's Stackelberg role. The next two terms show how the total demand for the brand name product is affected by a change in the share of informed consumers. Since informed consumers are aware of the cheaper generic substitutes, they buy less from the brand name firm than uninformed consumers and the sum of these terms will be negative. That is, the brand name firm, unlike the generic companies, loses sales if there is an increase in the share of informed consumers, which reduces the revenue loss from price reductions. Therefore, this works for a higher price reduction for brand name drugs compared to the generic drugs. The sign of the final term depends on the price change of the generic firms.

Given our assumptions, the theoretical model shows that it is likely that the substitution reform, which is our exogenous measure of increased consumer information about pharmaceutical prices, will lead to price reductions for brand name drugs while the predictions are less clear for the generic products. We now turn to the task of empirically testing the predictions of our model, i.e. that the prices of brand name (and perhaps also generic) pharmaceuticals will be reduced by the substitution reform.

4 The empirical analysis

4.1 Data and empirical specification

The pharmaceutical prices in Sweden during the period January 2001 to October 2006 are extracted from a dataset provided by the County Council of Västerbotten, Sweden. The data cover all prescription pharmaceuticals sold in the county during the period and sums up to a total number of 15 million prescriptions.

Only observations referring to pharmaceuticals considered to be exchangeable within the substitution reform are relevant for this study. We identified these observations by a two step procedure. First, we used a wide definition of which pharmaceuticals that should be considered to be exchangeable within the substitution reform. This wide definition is based on the three criteria specified by the Medical Product Agency, namely that the different pharmaceutical products have the same active substance, are of the same form (pills, oral fluids, etc.) and that the packages are of similar size. In a second step, these commodities were more carefully examined by manually comparing them to the Medical Product Agency list over substitutable products published at the time of the introduction of the reform. After this examination 4082 commodities, in 856 different exchange groups, are defined to be substitutes according to the regulations set out in the exchange reform. These commodities account for approximately 7 million prescriptions in our original sample. We can identify in which month each prescription is sold, and there are thus a maximum of 70 observations for each commodity, 21 before the reform and 49 after, giving a maximum number of observations of 285 740 (=4082*70). However, not all of these pharmaceutical products are sold in Västerbotten each month, reducing the actual number of observations to 126 904. Due to missing data for one or more of our key variables, the final sample consists of 105 587 observations.

Our dataset does not include information about which product that is the brand name product. From the theoretical model we know that brand name firms will distinguish themselves by setting higher prices than their competitors. This is also supported by previous empirical studies (e.g. Aronsson et al., 2001; Bond and Lean, 1977; Statman, 1981; Hurwitz and Caves, 1988). Thus, $Brand$ is an indicator variable that takes the value one if the pharmaceutical has had the highest price in the exchange group both in the first and third six months periods prior to the reform. The six month period is used to reduce the problem of not all commodities being sold each month, and two examination periods are used to reduce the risk of our labeling being affected by temporary price changes. We choose examination periods prior to the reform, since the reform may reduce price differences between brand name and generic drugs, making labeling based on periods after the reform less robust. In the 9 percent of the cases where the pharmaceutical had the highest price in only one of these periods, a closer examination was performed where a pharmaceutical was classified as a brand name drug if there were only one producer not belonging to the group of the most common generic producers in the Swedish market.[6] The remaining pharmaceuticals, where no apparent brand name producer could be identified were all treated as generics.[7]

As stated earlier, the main purpose of this paper is to study the effect of an increase in consumer information due to the introduction of the substitution reform on the prices of brand name and generic pharmaceuticals. Therefore, our focus will be to achieve unbiased estimates of the reform-effect. As such, we use product-specific fixed effects to capture several of the effects that according to equations (3) and (5) can explain the difference in price levels between pharmaceutical products, and focus our attention on changes in the prices caused by changes in the shares of informed consumers, as described in equations (6) and (7). The latter two equations reveal several variables which will influence the reform-effect. We will therefore take them as starting points to study heterogeneity in the reform-effect as well. Since we are not able to model all differences between the two equations, we estimate the models separately for brand name products and generics, allowing the parameter estimates to differ. However, the specification for the two groups will include the same variables. Below, we therefore only discuss the specifications for one of the two groups, brand name products.

From equation (7) we know that the effect of an increase in α on the brand

[6] In Sweden, there are some producers that specialize in producing and selling generic products in several of the substitution groups covered in this study. These are easily identified in our dataset.

[7] All regressions have also been performed excluding these 9 percent of the cases. All qualitative results from the estimations presented below are unchanged, and these results have thus been excluded in order to save space.

name price can be written

$$dP_b = \left(P_b - \frac{dC_b}{dQ_b}\right) * \left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g}\frac{\partial P_g}{\partial P_b} - \frac{dD_b^{UI}}{dP_b}\right)/(-\delta_b)d\alpha$$
$$+(D_b^I - D_b^{UI})/(-\delta_b)d\alpha + \frac{dP_b}{dP_g}\frac{dP_g}{d\alpha}/(-\delta_b)d\alpha. \quad (8)$$

The first product on the right hand side includes the following four terms: the markup over marginal cost, the difference in own price response of informed and uninformed consumers, the second derivative of the brand name firm's profit maximization problem, and the change in information due to the introduction of the substitution reform. Our data allow us to study heterogeneity in the first and forth terms.[8] First, the markup is assumed to be a linear function of our proxy variable $\ln markup$, that is

$$\left(P_b - \frac{dC_b}{dQ_b}\right) = \gamma_0 + \gamma_1 \ln markup_i. \quad (9)$$

$\ln markup_i$ is the logarithm of the difference between the average price for pharmaceutical product i in the pre-reform period and the minimum price in the group of exchangeable products during that period.[9] As such, this variable is defined for pharmaceutical i. Mathematical formulas used to calculate this and other proxy variables are presented in Appendix 3. Second, the change in information is written

$$d\alpha = \gamma_3 dD_t + \gamma_4 d(info_g * D_t), \quad (10)$$

where D is an indicator variable taking the value one after the introduction of the substitution reform, and $info_g$ is our proxy for the level of information available for consumers in exchange group g before the introduction of the reform. This is defined as the extra expenditure on pharmaceuticals in exchange group g prior to the reform compared to the expenditure if the given quantity would

[8] We do not model any heterogeneity in the second and third term. This reduces the number of interaction terms in our full model, but, similarly to the omission of interaction terms later discussed in Appendix 5, this do not cause any bias.

[9] Using the minimum price in the group of exchangeable products as a proxy for marginal cost is similar to the approaches adopted by Grabowski and Vernon (1992) and Rudholm (2001). The logarithm is used since the dependent variable will also be in logarithmic form. γ_0 is included to capture the average of the difference between the minimum price and the marginal cost.

have been bought at the lowest price in the exchange group each month, measured in percent. Following Sorensen (2000), this measure is based on the notion that if consumers systematically and over time pay more than necessary for a given pharmaceutical within an exchange group, this reflects that consumers in that group are less informed about prices of available generic substitutes than in other groups, cet. par. The idea is that the change in the share of informed consumers will be larger in markets with low initial information. Multiplication gives the following expression for the first product in equation (8)

$$\begin{aligned} \left(P_b - \frac{dC_b}{dQ_b}\right) * d\alpha &= \gamma_0\gamma_3 dD_t + \gamma_0\gamma_4 d(info_g * D_t) \\ &\quad +\gamma_1\gamma_3(\ln markup_i * dD_t) \\ &\quad +\gamma_1\gamma_4[\ln markup_i * d(info_g * D_t)]. \end{aligned} \tag{11}$$

In the second product in equation (8), our division of the sample in brand name and generic pharmaceuticals in part controls for differences in the term $D^I - D^{UI}$ between pharmaceutical firms, and we do not model any remaining heterogeneity in this term or in the second derivative, δ_b. What remains in the second product is $d\alpha$, which is measured as described above.

The third product in equation (8) shows that the change in the price of the brand name producer can also be affected by price changes of generic products, caused by a change in the proportion of informed consumers. The effect on the generic prices in an exchange group mostly depends on the same variables as the effect on the brand name prices. In addition it depends on the markup of the generic firms, which does not directly affect the price change of the brand name drug. To study whether this indirect effect has any significant effect on the brand name price, we therefore identify the term $\frac{dP_g}{d\alpha}$ by $diff \ln markup_i$, which is the difference between the average markup in the exchange group g and the markup for the brand name drug i.[10] $d\alpha$ is measured as presented above and we do not model any heterogeneity in the term $\frac{dP_b}{dP_g}$. Therefore, the third product in equation (8) leaves us with the two additional interaction terms $diff \ln markup * D$ and $diff \ln markup * info * D$.

So far, only the variables of main interest for this study as presented in equation (8) have been discussed, i.e. the variables that are supposed to capture

[10] This specification for the logarithmic markup of the other drugs in the exchange group is chosen in order to reduce the multicollinarity problem which otherwise could be a problem.

the effects of increased consumer information due to the substitution reform. However, as can be seen in equation (5), the price will also depend on n, which in our empirical model will be measured by the number of products sold in a specific substitution group each month. The interaction terms $n*D$ and $n*info*D$ are included since the studies of Stiglitz (1979) and Frank and Salkever (1992) indicated that the effect on prices of the number of products can be affected by the share of informed consumers. A trend variable, $Trend$, is included in order to account for possible common price trends. Finally, product-specific fixed effects, θ_i, have been included in order to allow for different intercepts for the I different products. The empirical specification of the most general price equation for brand name- and generic pharmaceuticals can thus be written

$$\begin{aligned}
\ln Price_{it} &= \beta_1 D_t + \beta_2(info_g * D_t) + \beta_3(\ln markup_i * D_t) \\
&\quad + \beta_4(\ln markup_i * info_g * D_t) + \beta_5(diff \ln markup_i * D_t) \\
&\quad + \beta_6(diff \ln markup_i * info_g * D_t) + \beta_7 n_{it} + \beta_8(n_{it} * D_t) \\
&\quad + \beta_9(n_{it} * info_g * D_t) + \beta_{10} Trend_t + \theta_i + \varepsilon_{it}. \qquad (12)
\end{aligned}$$

Descriptive statistics for all variables included in the empirical estimations are presented in Table 1. $\ln Price$ is the logarithm of the price of the pharmaceutical per one hundred defined daily dozes expressed in Swedish crowns in fixed October 2006 prices.[11] Comparing the descriptive statistics for brand name and generic products, prices are roughly the same in the two groups. However, as can be seen in Table 1, the markup over marginal cost is larger for the sample containing brand name pharmaceuticals. Also, the descriptive statistics for $diff \ln markup$ show that the competing products in an exchange group on average have a lower price than the brand name product, while the opposite is true for generics. As such, the similar prices for brands and generics in Table 1 are due to a selection effect, where generic firms to a large extent have chosen to enter markets where the price is high. In addition, Table 1 shows that the number of products is larger in the subsample for generics, as is the level of previous information.

[11] Defined daily doses (DDDs) is a World Health Organization measure of drug quantity.

Table 1. Descriptive statistics

	Brands		Generics	
Variable	Mean	Std. Dev.	Mean	Std. Dev.
$Price$	1796.04	4359.74	1727.62	5137.14
$\ln Price$	6.64	1.18	6.39	1.31
D	0.67	0 .47	0.67	0 .47
$info$	7.98	22.51	12.90	81.63
$\ln markup$	4.34	1.49	3.21	1.62
$diff \ln markup$	-0.28	0.59	0.97	1.12
n	2.35	1.44	3.29	1.56
$Trend$	33.99	20.05	33.09	19.45
$info * D$	5.15	16.44	8.74	65.17
$\ln markup * D$	2.90	2.37	2.10	1.99
$\ln markup * info * D$	26.46	101.03	32.65	371.04
$diff \ln markup * D$	-0.18	0.51	0.69	1.08
$diff \ln markup * info * D$	-1.10	9.34	13.91	99.60
$n * D$	1.66	1.70	2.24	2.07
$n * info * D$	15.02	55.61	28.86	119.17
Nr. $\ln markup$	47008		57861	
Nr.	47121		58466	
Products	873		1613	

Nr. $\ln markup$ is the number of observation for the variables including $\ln markup$ or $diff \ln markup$ while Nr. is the number of observations for all other variables.

In equation (12), the following variables might be endogenous in the sense that these variables correlate with the error term; $info$, $\ln markup$, $diff \ln markup$ and n. The available dataset does not contain any reasonable variables to be used as instruments, hence endogeneity has to be addressed in some other way. In order to be able to later conclude which result that might be driven by endogeneity, we therefore discuss in Appendix 4 how the correlation between the error term and these variables are expected to affect the estimates of the different parameters. One should however keep in mind that these variables are proxy-variables and that the estimators can therefore also be biased due to measurement errors. Since bias in the estimators related to the variables mentioned above might affect the estimated reform-effect, we show in Appendix 5 how we can create a simple model, without the interaction variables, in order to estimate the reform-effect. In short, we show that the estimator for the reform parameter, β_1, will be an unbiased estimator of the average reform-effect if the interaction variables are left out from the specification. The explanation is that this estimator will capture the linear relationship between this variable and the omitted interaction variables (e.g. Greene, 2003, Chapter 8).

Due to low variation in the number of products during the period of study, the remaining possibly endogenous variable, n, most likely only has a very limited influence on the estimated reform-effect. During the study-period, it is likely that the bulk of the correlation between n and the reform is due to the effect the reform has on the number of products, and that only a smaller fraction of the correlation is due to exogenous variation in n. Therefore, not controlling for n in the regression probably gives a better estimate of the total reform-effect.[12] As such, the specification which will be used to estimate the average effect of the reform for both brand name- and generic pharmaceutical subsamples is written

$$\ln Price_{it} = \beta_1 D_t + \beta_{10} Trend_t + \theta_i + \varepsilon_{it}, \tag{13}$$

where all included variables are exogenous. In all estimations, we use a Prais-Winsten estimator which corrects for first order serial correlation in the error terms. In addition the error terms are allowed to be heteroskedastic and correlated within ATC-groups.[13]

4.2 The adjustment process

Two circumstances give firms an incentive to gradually adjust their price after the reform.[14] First, there is incomplete information for firms about the reactions of consumers and other firms to the reform. Second, price increases are rarely

[12] We will also estimate the basic model including the number of firms, n, in order to study how this affects the parameter estimates for the reform indicator variable.

[13] In the World Health Organization's Anatomical Therapeutic Chemical (ATC) classification system, the drugs are divided into different groups according to the organ or system on which they act, and their chemical, pharmacological and therapeutic properties. In the ATC-groups used here, drugs which share the same chemical substances are grouped together.

[14] Since pharmaceutical firms knew about the reform before it came into effect in October, 2002, it is also possible that they started to adjust to the reform before this date. If this was the case, we would expect to obtain larger estimates of the reform effect if the estimations were performed excluding observations from the months directly before the reform. To study

allowed by the PBB. To capture this, we introduce the variable $D/(t-R)$, where R is the time for the reform. We let the denominator be raised to the power μ, where μ is a parameter that will measure the curvature of the adjustment process. As such, our basic empirical specification including the adjustment process is written

$$\ln Price_{it} = \beta_1 D_t + \beta_{10} Trend_t + \gamma[D_t/(t-R)^{\mu}] + \theta_i + \varepsilon_{it}. \qquad (14)$$

The models with adjustment are non-linear in the adjustment variable $D/(t-R)$. Since the models are nonlinear only due to one parameter, μ, it is convenient to estimate the models using a grid-search estimation strategy. We have employed this method for each model setting μ to values ranging from 0 to 4 and then estimating the remaining parameters using a Prais-Winsten estimator. Finally, likelihood values were used to discriminate between the different parameter values. The likelihood values were as well used to calculate 95 percent confidence intervals for the adjustment parameter, μ. As can be seen in the notes to Tables 2 and 3, the confidence intervals are not symmetric around the point estimates. This is expected since a value of μ equaling zero lead to an empirical model where the adjustment variable becomes equal to the reform indicator variable.

4.3 Results

The results from the estimation for brand name products are presented in Table 2. In order to study heterogeneity in the reform-effect, we first discuss the results of the full model. For all variables that interact with other variables, we only discuss the differential and the derivatives, respectively, which are evaluated at the mean value for each variable. The estimate of the differential $\Delta LnPrice/\Delta D$ from the full model shows that the average reduction in the price of brand name pharmaceuticals due to the reform is approximately 2 per-

the importance of this possibility, we estimated all models excluding observations originating from April 2002, when the law regarding the reform was passed by parliament, until October 2002, as well as from January 2002, when the bill was presented to parliament, until October 2002, respectively. This did not change the estimated reform effect significantly in any of the estimated models, and in the majority of the cases the estimated reform effect actually became slightly smaller. Thus, these results indicate that the potential adjustments before the reform were too small to have any important impact on our estimates.

cent,[15] and there is a significant negative trend showing an average decrease in pharmaceutical prices of 0.2 percent per month. From the theoretical model it is expected that if consumers had a low level of information prior to the reform (i.e. they systematically spent more money than necessary on the pharmaceuticals they bought), the reduction in brand name price after the introduction would be large. As can be seen in the lower half of Table 2, this is also the case. However, this result must be interpreted with caution since the estimators of the parameters for $info$ may be negatively biased, as discussed in Appendix 4.

The point estimate of the interaction between the reform and $\ln markup$ in Table 2 shows that the effect of the reform was more pronounced if the markup was large. However, this result is not statistically significant at conventional levels. Neither does the indirect effect from equation (8) (as proxied by $diff \ln markup$ in our empirical model) show any significant impact on the reform-effect on pharmaceutical prices. The positive point estimate for this variable may be explained by a positive correlation between this term and the error terms. Finally, expanding the informed section of the market through the introduction of the reform is found to increase the impact of the number of products on pharmaceutical prices, while the number of products had no significant effect on the prices themselves.

We estimate the basic model with and without n in order to study if the inclusion of this variable affects the estimate of the reform-effect. As can be seen in Table 2, the parameter estimate for the number of products is not statistically significant, and the estimated reform-effect is similar in size in both models. As such, the reform-effect does not seem to be affected by an exclusion of the number of products from the empirical model. The estimated reform-effect in these two models is slightly lower than in the full model, but still approximately 2 percent. There is also a negative time trend of 0.2 percent per month.

[15] Since the dependent variable is in logarithmic form, the exact change in price (in percent) should for dummy variables be calculated using the formula $100 * [\exp(\beta) - 1]$. However, since the parameter estimates are small, using the exact formula gives the same results as using the parameter estimates directly after rounding.

Table 2. Brand name prices

	Full	Basic+n	Basic	Basic+adj.
D	22.32**	-18.85***	-18.86***	-56.88***
	(11.13)	(5.71)	(5.70)	(17.19)
$Trend$	-2.01***	-2.01***	-2.01***	-1.53***
	(0.51)	(0.52)	(0.50)	(0.47)
n	1.60	-0.23		
	(1.36)	(0.79)		
$info * D$	-4.35*			
	(2.48)			
$\ln markup * D$	-6.25**			
	(2.94)			
$\ln markup * info * D$	0.71*			
	(0.39)			
$diff \ln markup * D$	1.98			
	(5.84)			
$diff \ln markup * info * D$	1.20*			
	(0.67)			
$n * D$	-2.28			
	(1.68)			
$n * info * D$	-0.04			
	(0.05)			
$D/(t-R)$ (γ)				38.14***
				(13.89)
$D/(t-R)$ (μ)				354.00
				(see note)
$\Delta \ln Price/\Delta D$	-19.25***	-18.85***	-18.86***	-42.40***
	(5.60)	(5.71)	(5.70)	(12.25)
$d \ln Price/d(info * D)$	-1.73*			
	(0.99)			
$d \ln Price/d(\ln markup * D)$	-0.62			
	(4.12)			
$d \ln Price/$	11.59			
$d(diff \ln markup * D)$	(7.37)			
$d \ln Price/d(n * D)$	-2.62*			
	(1.54)			
$d \ln Price/dn$	-0.15	-0.23		
	(0.77)	(0.79)		
Sample size	47008	47121	47121	47121
Log likelihood	76724	76858	76859	76868

The reported values are the estimates multiplied by 1000. The regressions include product-specific fixed effects. Robust standard errors, conditioned on the parameter, μ, are reported in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. The 95%-confidence interval for μ is $0<\mu<1086$. The differential/derivatives are evaluated at the mean for each variable. For $D/(t\text{-}R)^{\mu}$ the mean in the post-reform period is used.

The results from the model with the adjustment variable show that the estimated reform-effect is larger than in the basic model, with an estimated average price reduction during the study-period of 4 percent. One explanation for this result is that including the adjustment parameter in the empirical specification changes the parameter estimate for the time trend from -0.20 to -0.15 percent per month. The estimate for the reform dummy in this model, minus 6 percent, can be interpreted as an estimate of the long run effect of the reform. The relatively large standard errors of this estimate, nearly 2 percent, is likely caused by that the correlation between D and $D/(t-R)^{\mu}$ is as high as 0.82 (e.g. Greene, 2003, Chapter 4). According to a likelihood-ratio test, we can reject the hypothesis of an instant adjustment at the time of the reform. We therefore regard the estimates from the model with the adjustment variable as the more reliable ones. The estimated reform-effect for the model with adjustment is illustrated in Figure 1.

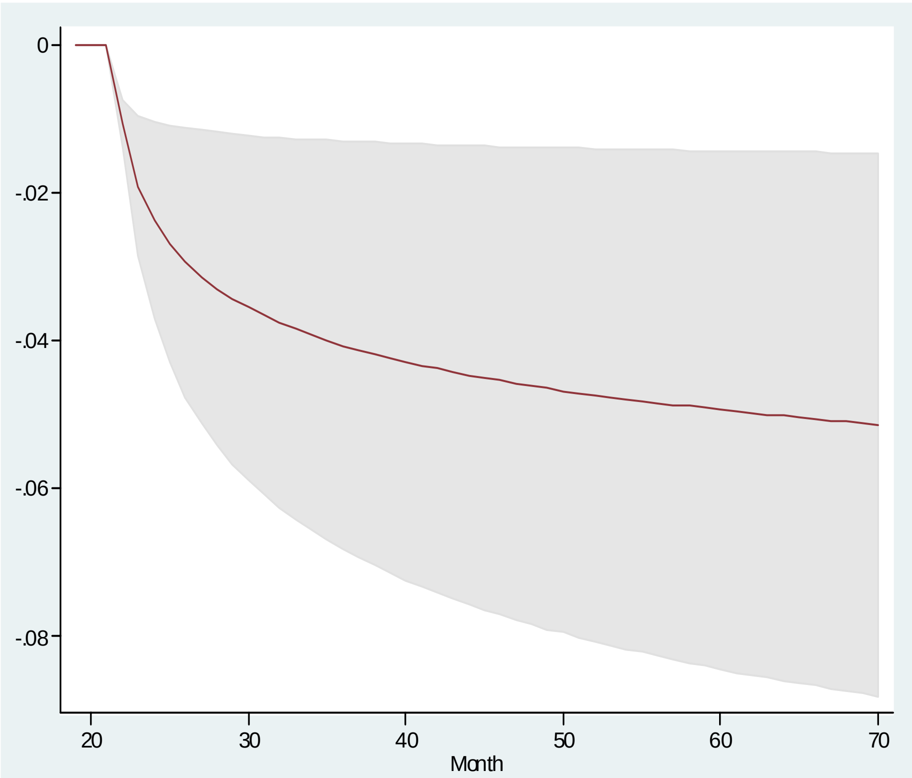

Figure 1. Brand name prices, estimated reform effect with 95% CI

Table 3. Generic prices

	Full	Basic+n	Basic	Basic+adj.
D	35.85***	-11.02***	-10.98***	-135.85**
	(12.08)	(1.70)	(1.70)	(55.49)
$Trend$	-2.67***	-2.75***	-2.74***	-2.01***
	(0.75)	(0.76)	(0.76)	(0.62)
n	2.39***	0.75		
	(0.73)	(0.57)		
$info * D$	-1.37			
	(1.03)			
$\ln markup * D$	-8.03***			
	(2.38)			
$\ln markup * info * D$	0.22			
	(0.17)			
$diff \ln markup * D$	-10.82***			
	(4.16)			
$diff \ln markup * info * D$	0.19			
	(0.14)			
$n * D$	-2.50**			
	(1.05)			
$n * info * D$	0.01			
	(0.01)			
$D/(t-R)$ (γ)				125.20**
				(54.51)
$D/(t-R)$ (μ)				101.00
				(see note)
$\Delta \ln Price/\Delta D$	-11.19***	-11.02***	-10.98***	-41.34***
	(1.79)	(1.70)	(1.70)	(14.38)
$d \ln Price/d(info * D)$	-0.44			
	(0.36)			
$d \ln Price/d(\ln markup * D)$	-5.13**			
	(2.53)			
$d \ln Price/$	-8.40**			
$d(diff \ln markup * D)$	(4.09)			
$d \ln Price/d(n * D)$	-2.41**			
	(1.01)			
$d \ln Price/dn$	0.77	0.75		
	(0.58)	(0.57)		
Sample size	57861	58466	58466	58466
Log likelihood	79506	79847	79843	79855

The reported values are the estimates multiplied by 1000. The regressions include product-specific fixed effects. Robust standard errors, conditioned on the parameter, μ, are reported in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. The 95%-confidence interval for μ is $0<\mu<672$. The differential/derivatives are evaluated at the mean for each variable. For $D/(t\text{-}R)^{\mu}$ the mean in the post-reform period is used.

The results concerning generic pharmaceutical prices presented in Table 3 indicate an average price reduction due to the reform of one percent in all three estimated models not including the adjustment variable. Further the estimates show a negative price trend of 0.3 percent per month. As such, the reform-effect without adjustment is smaller in size for generics than for brand name pharmaceuticals, while the price trend is larger. Neither of the differences is statistically significant.

The point estimates of the derivatives at the bottom of Table 3 show that both high values of $info$ and $\ln markup$ enhance the effect of the reform as expected, but the effect of $info$ is not statistically significant at conventional levels. For generics, the parameter estimate for $diff \ln markup * D$, is negative and statistically significant at the five percent level. Our interpretation of this is that a high markup prior to the reform of the competitors to generic i increases the reform-effect on the competitors, which in turns increases the need for generic i to lower its price. As discussed in Appendix 4, potential endogeneity will cause the estimators related to this variable to be positively biased. As such, this gives further support that the results are in line with our theoretical model. Finally, as in the estimations concerning brand name pharmaceuticals, expanding the informed section of the market through the introduction of the reform is found to increase the impact of the number of products on pharmaceutical prices, but the number of products had no significant effect on the prices themselves.

Turning to the results from our model including the adjustment process, the results show that the reform-effect is larger than in the basic models, with an estimated average price reduction of 4 percent. As for brand name pharmaceuticals, a large part of the difference between the models can be explained by the difference in the parameter estimates for the time trend. The point estimates for the parameter μ, indicate that the adjustment path for generics is less curved than that for brands. However the divergence is not statistically significant different from zero. Further, the low value of μ results in a correlation between D and $D/(t-R)^{\mu}$ of 0.99, which increases the standard errors for the parameter estimates related to these variables (e.g. Greene, 2003, Chapter 4). The standard errors for the two estimates are over 5 percent and it is therefore not meaningful to discuss these estimates. However, the standard error for the total reform-effect is not influenced by the high correlation between D and

$D/(t-R)^{\mu}$. As for the brand name sample, we can reject the hypothesis of an instant adjustment at the time of the reform, and we therefore regard the estimates from this model as the most reliable ones. The estimated reform-effect for the model with adjustment is illustrated in Figure 2.

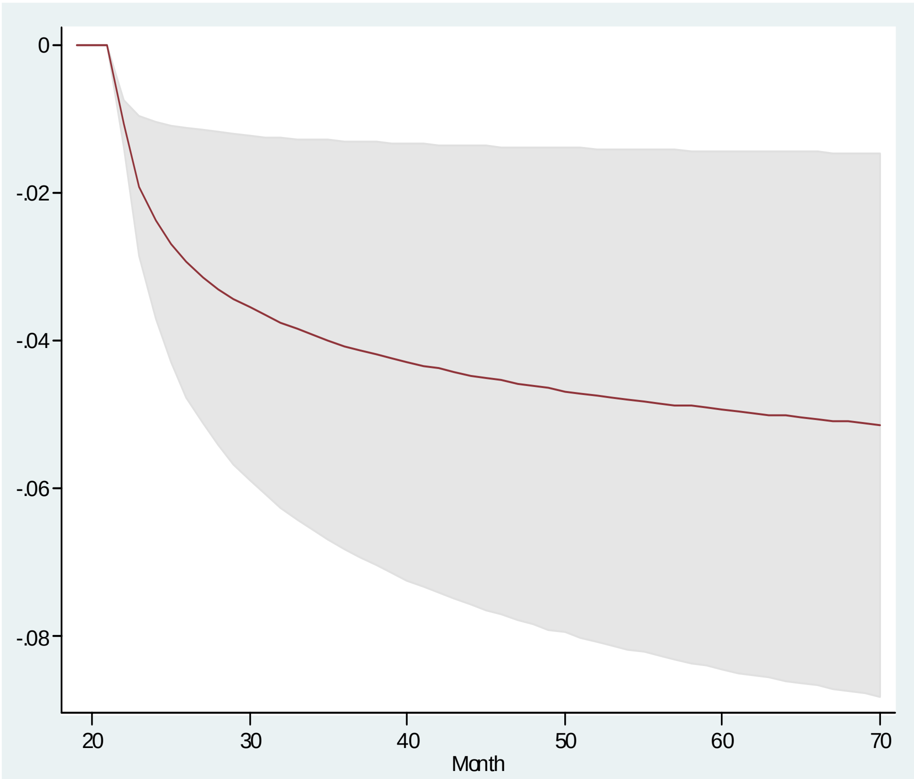

Figure 2. Generic prices, estimated reform effect with 95% CI

Taken together, the results indicate that the effects of the reform on the price paths of pharmaceutical products are more similar for brand and generic products than expected, with an average price reduction due to the implementation of the reform of about 4 percent. These results could be compared to previous studies concerning the effects of substitution reforms on pharmaceutical prices (e.g. Buzzelli et al., 2006). In comparison, their results show an average price reduction of 3 percent after the implementation of the substitution reforms. It should, however, be noted that they use pharmaceutical price index data, from 16 OECD countries, and that their estimate therefore is not directly comparable to ours. Firstly, their results include the effect that the substitution reforms have had on the average price through a change in the composition of drugs. Secondly, their estimate is an average for all drugs, not only those directly affected by the reforms. Further, the study by Buzzelli et al. does not incorporate any adjustment process in the empirical specification.

5 Discussion

As noted by Stiglitz (1979), it has long been recognized that imperfect consumer information give firms some degree of monopoly power which leads to prices above competitive levels. In addition, Salop and Stiglitz (1977) showed that imperfect information could lead to market equilibria where homogenous products are sold at different prices. The situation in the Swedish pharmaceuticals market before the introduction of the exchange reform was characterized by both these phenomena, since close substitutes such as brand name- and generic pharmaceutical products were sold at different prices, and since especially brand name products seemed to be priced well above long run marginal costs.

In this paper, the introduction of the Swedish substitution reform was used as a natural experiment regarding the effects of increased consumer information on brand name and generic pharmaceutical prices. The main hypothesis to be tested was if increased consumer information about pharmaceutical prices and available generic substitutes due to the introduction of the substitution reform had led to any reduction in the price of brand name and/or generic pharmaceuticals. Another hypothesis to be tested was whether the possible price response due to the introduction of the reform differed between brand name- and generic drugs.

First, following Frank and Salkever (1992), a theoretical model was set up and the effects of increased consumer information on pharmaceutical prices were analyzed. The model indicated that an increase in information would, under reasonable assumptions, led to a price reduction for brand name products, while the results for generic pharmaceuticals were more ambiguous. Based on our model, we also expected that the reduction in price would be large if the markup over marginal cost was large and/or the market had been characterized by low levels of information before the reform.

The results form the empirical part of the paper show an average reduction in prices due to the reform of about 4 percent during the study-period, both for brand name- and generic pharmaceuticals. In addition, the results give some support for the reform-effect being amplified for pharmaceuticals in markets which had previously been characterized by low levels of consumer information, as well as for pharmaceuticals which prior to the reform had high markups over marginal cost.

We are also able to present tentative results regarding how an increase in information will affect the impact of the number of products on pharmaceutical prices. The results presented in this paper are in line with Grabowski and Vernon's (1992) conjecture, that expanding the informed portion of the market should increase the price lowering effect of entry.

A final point that needs to be discussed is how the brand name product has previously been able to set a high price relative to the generic substitutes without losing market share in the absence of the substitution reform. Stiglitz (1979, p 340) suggests that "a flow of ignorance can be maintained either by entry of new firms or new individuals". This could perhaps make it possible for, for example, a brand name producer to price its product high relative to its generic substitutes without losing market shares. To empirically investigate this question is outside the scope of the present paper, but warrants future research.

Appendix 1. Indirect effects

The effect on the price of generic i of a change in the price of the other generics is written

$$\frac{dP_i}{dP_j} = \left[\left(P_i - \frac{dC_i}{dQ_i} \right) \alpha \frac{d^2 D_i^I}{dP_i dP_j} + \alpha \frac{dD_i^I}{dP_j} \right] / (-\delta_i).$$

The first term includes $d^2 D_i^I / dP_i dP_j$, which describes how the slope of the demand curve facing firm i is affected by a change in the price of other generics. This term will be negative if a price reduction of other drugs changes the marginal consumer to one who is less price sensitive than the previous one. However, without knowing the specific form of the demand function, this term cannot be signed. Since the drugs are substitutes, the second term is positive and states that if the other prices are reduced, this will lower the demand for drug i, which works for a price reduction of drug i. The price response of firm i to a price change by the brand name firm and the price response of the brand name firm to a price change by the generic ones can be written correspondingly. Theoretically, we cannot rule out the possibility that the first term is negative and dominates the second one without making more explicit assumptions about the demand function. In this paper we will assume that $\frac{dP_i}{dP_j} > 0$, $\frac{dP_i}{dP_b} > 0$ and $\frac{dP_b}{dP_g} > 0$, without assuming a specific demand function. Among the possible demand functions that would result in the positive derivatives that we assume, a simple linear demand function like $D_i^I = d_1 - d_2 P_i + d_3 P_j + d_4 P_b$ can be mentioned.

Appendix 2. Second order conditions

The second order profit maximization condition for generic firm i, is written

$$\begin{aligned} d^2\pi_i / dP_i^2 &= \delta_i = \left(P_i - \frac{dC_i}{dQ_i} \right) * \left[\alpha \frac{d^2 D_i^I}{dP_i^2} + (1-\alpha) \frac{d^2 D_i^{UI}}{dP_i^2} \right] \\ &\quad +2 * \left[\alpha \frac{dD_i^I}{dP_i} + (1-\alpha) \frac{dD_i^{UI}}{dP_i} \right] < 0. \end{aligned}$$

The corresponding expression for the brand name firm is written

$$\begin{aligned} d^2\pi_b/dP_b^2 &= \delta_b = \left(P_b - \frac{dC_b}{dQ_b}\right) * \\ & \left[\alpha\left(\frac{\partial^2 D_b^I}{\partial P_b^2} + 2\frac{\partial^2 D_b^I}{\partial P_g^2}\left(\frac{\partial P_g}{\partial P_b}\right)^2 + \frac{\partial D_b^I}{\partial P_g}\frac{\partial^2 P_g}{\partial P_b^2}\right) + (1-\alpha)\frac{dD_b^{UI}}{dP_b}\right] \\ & +2 * \left[\alpha\left(\frac{\partial D_b^I}{\partial P_b} + \frac{\partial D_b^I}{\partial P_g}\frac{\partial P_g}{\partial P_b}\right) + (1-\alpha)\frac{dD_b^{UI}}{dP_b}\right] < 0. \end{aligned}$$

Appendix 3. Definitions of empirical measures

Denote the lowest observed price in exchange group g prior to the reform by $MinPrice_g$. Let $Sales_{it}$ denote the sales of pharmaceutical i at period t and let $Sales_{gt}$ be the corresponding for all pharmaceuticals in an exchange group. Then $info_g, \ln markup_i$ and $diff \ln markup_i$ is written as follows

$$\begin{aligned} info_g &= \left[\sum_{t=1}^{R}\left(\frac{\sum_{i=1}^{n} Price_{it} * W_{it}^1}{MinPrice_{gt}}\right) * W_{gt}^1 - 1\right] * 100 \\ \ln markup_i &= \ln\left[\sum_{t=1}^{R} Price_{it} * W_{it}^2 - MinPrice_g\right] \\ diff \ln markup_i &= \ln\left[\sum_{i=1}^{n}\sum_{t=1}^{R} Price_{it} * W_{it}^3 - MinPrice_g\right] \\ & - \ln markup_i, \end{aligned}$$

where

$$\begin{aligned} W_{it}^1 &= Sales_{it}/Sales_{gt}; \qquad W_{gt}^1 = Sales_{gt}/\left(\sum_{t=1}^{R} Sales_{gt}\right); \\ W_{it}^2 &= Sales_{it}/\left(\sum_{t=1}^{R} Sales_{it}\right); \qquad W_{it}^3 = Sales_{it}/\left(\sum_{t=1}^{R} Sales_{gt}\right). \end{aligned}$$

That is, $info_g$ is defined as the extra expenditure on pharmaceuticals in exchange group g prior to the reform compared to the expenditure if the given quantity would have been bought at the lowest price in the exchange group each month, measured in percent. $\ln markup_i$ is measured as the logarithm of the difference between the average price for pharmaceutical product i in the pre-reform period and the minimum price in the group of exchangeable products during that period. Finally, $diff \ln markup_i$ is created by subtracting $\ln markup_i$ from the logarithm of the difference between the average and the lowest price in one exchange group in the pre-reform period.

Appendix 4. Endogeneity

By recalling the definition of the variable $info$ from Appendix 3, we see that this variable depends on prices and therefore could be correlated with the error term of equation (12). For brand name pharmaceuticals, an increase in the pre-reform error terms will lead to an increase in $info$. The inclusion of product-specific fixed effects operates so that the sum of the error terms for a pharmaceutical during the whole study-period approaches zero. Therefore an increase in the pre-reform error terms for a pharmaceutical will lead to a reduction in the post-reform error terms for that pharmaceutical. For brand name firms, this results in a negative correlation between the post-reform error terms and the variable $info$, which will lead to negative bias in the estimators related to $info$. (The correlation between $info$ and the pre-reform error terms does not directly affect the bias, since $info$ is only included in the model together with the reform dummy.)

For the cheapest generic in an exchange group, an increase in the pre-reform error terms will lead to a reduction in $info$, since the denominator will increase proportionally more than the numerator. Therefore, the post-reform error term of this pharmaceutical will be positively correlated with $info$. This is also true for a uniform increase of the pre-reform error terms for all generics in an exchange group. In the estimation for generic pharmaceuticals, the estimators related to $info$ are therefore likely to be positively biased.

All pre-reform error terms for a pharmaceutical have a positive effect on the variable $\ln markup$, except the error term for the observation that affect $MinPrice$ which has a negative effect. Since in each exchange group there is only one observation of $MinPrice$, and several observations of other pre-reform prices, it is reasonable to expect the correlation between the lowest pre-reform price and the post-reform error terms to be smaller in absolute size than the correlation between all pre-reform prices and these error terms. The variable

$\ln markup$ is therefore expected to be negatively correlated with the post-reform error terms, making the estimators related to this variable negatively biased. Since an increase in $\ln markup$ will lead to a decrease in $diff \ln markup$, the estimators related to the latter variable are positively biased.

The variable n is likely to be endogenous as well. An increase in the size of the error term will, through its effect on the price, lead to an increase in the number of products. This will, in turn, lead to positive bias in the estimators related to n. Including n in the estimation and treating it as exogenous might also lead to a negative bias in the estimator of the reform-effect. The reason is that the reform is likely to have a negative effect on prices, which in turn could have a negative effect on the number of products. If the number of products is estimated to have a negative effect on the prices, conditioning on it will then work for an overestimation of the negative price effect of the reform. However, in addition the reform makes it more likely that a generic product will be dispensed, increasing the probability of generic entry. Including n could therefore also lead to an underestimation of the negative price effect of the reform.

Appendix 5. Measuring the reform-effect

Denote the reform indicator variable by the vector $\mathbf{D}$. Let $\mathbf{X}(\mathbf{D})$ denote the $it{*}7$ matrix for the interaction variables $(info * D)$, $(\ln markup * D)$, $(\ln markup * info * D)$, $(diff \ln markup * D)$, $(diff \ln markup * info * D)$, $(n * D)$ and $(n * info * D)$ and let $\boldsymbol{\gamma}$ be a vector of parameters related to these variables. Ignore for simplicity that our model also includes the variables n and $Trend$ and product-specific fixed effects, θ_i. Then our structural model is written

$$\ln \mathbf{Price} = \mathbf{X}(\mathbf{D})\boldsymbol{\gamma} + \mathbf{D}\beta_1 + \boldsymbol{\varepsilon}.$$

If this equation is estimated on reduced form, excluding $\mathbf{X}(\mathbf{D})$, the estimator for β_1 becomes according to Greene (2003, Chapter 8).

$$b_1 = \beta_1 + (\mathbf{D}'\mathbf{D})^{-1}\mathbf{D}'\mathbf{X}\boldsymbol{\gamma} + (\mathbf{D}'\mathbf{D})^{-1}\mathbf{D}'\boldsymbol{\varepsilon}.$$

Now, in order to simplify things assume that there are observations in only two periods, one before the reform ($t = 1$) and one after ($t = 2$), and let us calculate what b_1 represents if we omit the interaction terms from our empirical specification. The matrices is written

$$\mathbf{D}'=\begin{pmatrix} \mathbf{0} & \mathbf{1} \end{pmatrix};\mathbf{D}=\begin{pmatrix} \mathbf{0} \\ \mathbf{1} \end{pmatrix};\mathbf{X}=\begin{pmatrix} \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{x}_1 & \mathbf{x}_2 & \mathbf{x}_3 & \mathbf{x}_4 & \mathbf{x}_5 & \mathbf{x}_6 & \mathbf{x}_7 \end{pmatrix};\boldsymbol{\varepsilon}=\begin{pmatrix} \varepsilon_1 \\ \varepsilon_2 \end{pmatrix},$$

where $\mathbf{0}$ and $\boldsymbol{\varepsilon}_1$ are vectors of zeros and error terms for the I pharmaceuticals before the reform, and where $\mathbf{1}$, $\mathbf{x}_1$, $\mathbf{x}_2...\mathbf{x}_7$ and $\boldsymbol{\varepsilon}_2$ are vectors of ones, the interaction variables and the error terms for the I pharmaceuticals after the reform. Matrix multiplication then gives

$$b_1=\beta_1+\sum_{k=1}^{7}\frac{1}{I}\sum_{i=1}^{I}x_{ki}\gamma_k+\frac{1}{I}\sum_{i=1}^{I}\varepsilon_{i2}.$$

That is, if we exclude the interaction terms with the reform dummy, the average effect of these variables will still be captured by the parameter estimate of the reform indicator variable.

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